Essays on globalization, monetary policy and financial crisis'

Qian, Z.

Publication date:
2012

Document Version
Publisher's PDF, also known as Version of record

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September 21, 2012

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Proefschrift

ter verkrijging van de graad van doctor aan Tilburg University op gezag van de rector magnificus, prof. dr. Ph. Eijlander, in het openbaar te verdedigen ten overstaan van een door het college voor promoties aangewezen commissie in de aula van de Universiteit op vrijdag 21 september 2012 om 10.15 uur door

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geboren op 28 maart 1983 te Chongqing, P. R. China.

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Acknowledgements

Over last three years, I have been working on my thesis at Tilburg University. Many people have helped me in this process. The first I would like to thank are my supervisors Hans Blommestein and Sylvester Eijffinger.

Hans was the second reader of my MPhil thesis which serves as a starting point for the fourth chapter of the PhD thesis. I benefit a lot from his comments on developing the fourth chapter. He is the coauthor of chapter 2 and 3 of the thesis. His knowledge about the financial market and institutions is extremely helpful for establishing valid assumptions for the models and understanding the results of the models. He arranged a trainee position for me at the Bond Market and Public Debt Management Unit in the Organisation for Economic Co-operation and Development (OECD) in the summer of 2011. At that time, the unit is doing research on the European sovereign debt crisis. My work there motivated the chapter on European sovereign credit default swaps.

Sylvester is the supervisor for both my MPhil and PhD thesis. I start to work with him since the beginning of my study in Tilburg. My favorite advice from him is that research must be fun. Together with him, I am able to study three interesting topics. Those studies produce the main chapters (2, 3, 4) of the thesis. Since he worked on related topics before, discussion with him improved my understanding of the topics. I worked as Sylvester's teaching assistant for the course seminar financial economics. I learned a lot from him on how to design and teach an interactive economics course. He also supported my visits to the Kiel Institute for the World Economy and the OECD, which helped me improve my research skills. Needless to say, without the support from the supervisors, it is impossible for me to find a nice research job close to home.

I want to thank Henk van Gemert, Jakob de Haan, Lex Hoogduin, and Jenny Ligthart for reading my manuscript and joining the PhD committee. Comments from the committee members are very helpful. Special thanks to Jenny for reading the chapters so carefully given her physical condition. I am also grateful for her study advices as my education coordinator and her support for my job applications as one of the recommenders.

I am grateful to Klaus Desmet, Benedikt Goderis, Kan Ji, Kebin Ma, Rob Nijskens, Peter van Oudheusden, Maria Fabiana Penas, Damjan Pfajfar, Louis Raes, Sjak Smulders, Roberto Rigobon, Harald Uhlig, Burak Uras, Gonzaque Vannoorenberghe, Wendun Wang, Huaxiang Yin and seminar participants at Tilburg University, the ENTER Jamboree (Toulouse School of Economics), the 6th Eurostat Colloquium on Modern Tools for Business Cycle Analysis for helpful discussion on preliminary versions of the thesis chapters. I thank Chang-Jin Kim who kindly provided me the Gauss code for his two-step regime switching model. I am also grateful to Davide Romelli for sharing the dynamic central bank independence data with me. Christoph Schottmüller provided a latex template for the PhD thesis. By using this template, I saved a lot of time.

I thank Benedikt Goderis and Sjak Smulders for writing recommendation letters for my job searching. Johannes Binswanger, Patricio Dalton, and Gonzaque Vannoorenberghe gave me a nice practice interview. Katie Carman provided useful comments on my job application package. Cecile de Bruijn helped sending the application packages. With their help, I managed to find a nice job at an early stage. This saved my time and energy for the last PhD research project.

I also want to express my gratitude to my friends and fellow students who made my life in Tilburg more exciting. Thanks the organizers of the Chinese and non-Chinese badminton clubs who provided my favorite way to relax from the research work. Thanks the secretaries of CentER, the department of Economics, the European Banking Center, and the Netherlands Network of Economics (NAKE) whose work helped my study and work.

Finally, I want to thank my family for supporting my study abroad. I owe a lot to my wife, Qian Luo. To support my academic career, she took all my family obligations. I cannot finish the thesis without such a generous support. My deepest gratitude for her love.

Contents

Chapter 1

Introduction

The subprime mortgage crisis and the ongoing European debt crisis have led macroeconomists to rethink macroeconomics (Caballero, 2010; Stiglitz, 2011). Standard macroeconomic models and their empirical counterparts put aside financial crises. Typical assumptions of those models, such as rational expectations, representative agents, complete financial market, limit their ability to analyze causes and consequences of financial crises. Chapter 2 and 3 of this thesis focus on the interaction between macroeconomic variables and the financial sector. They relax assumptions of the standard macroeconomic models in different directions. Chapter 2 relax the rational expectations assumption. Chapter 3 relax the assumption of representative agents. Both chapters relax the complete financial market assumption.

Chapter 2 studies the determinants of the sovereign credit default swap (CDS) spreads of five Euro-area countries (Greece, Ireland, Italy, Portugal and Spain). We focus on the period in which the global financial crisis deepens and contingent government debt necessary to bailout the financial sector may have contributed to the run-up to sovereign debt crises in those countries. The sovereign CDS contract is a quasi-insurance instrument for the sovereign credit risk. In the previous literature, it is believed that macroeconomic fundamentals linked to a country's sovereign credit risk should affect the price of this quasi-insurance instrument, that is, the sovereign CDS spread. We find that there are regime switches in the process of sovereign CDS spread changes. Under one regime, changes in financial market-based indicators of macroeconomic fundamentals have significant explanatory power to changes in sovereign CDS spreads. Under the other regime, changes in financial market-based indicators of macroeconomic fundamentals have no explanatory power to changes in sovereign CDS spreads. Those regime switches are difficult to explain under the rational expectations assumption. By contrast, they are consistent with a theory of "animal spirits" in which agents are cognitively limited.

The concept "animal spirits" dates back to Keynes (1936). He argues that if market uncertainty is high, an investor's decision is "the result of animal spirits-a spontaneous urge to action rather than inaction, and not as the outcome of a weighted average of quantitative benefits multiplied by quantitative probabilities". Although animal spirits play an important role in the economic theory of Keynes (1936), it is not at the core of modern macroeconomics.[1] On the occasions when the term "animal spirits" is used in macroeconomic models, it is interpreted as sunspot shocks to the expectations of rational investors who are not cognitively limited.[2] Chapter 2 shows that the rational expectations interpretation of animal spirits cannot explain what we find in the European sovereign CDS market. However, an alternative interpretation of animal spirits provides a good explanation for the sovereign CDS spread dynamics we observe. More specifically, animal spirits is interpreted as the belief of cognitively limited agents. If market uncertainty is low, cognitive biases are small and market-based indicators of macroeconomic fundamentals contain useful information for investors. Therefore, those indicators affect sovereign CDS pricing. If market uncertainty is high, cognitive biases are large and market-based indicators of macroeconomic fundamentals become useless for investors to infer the sovereign credit risk. Therefore, those indicators no longer affect the sovereign CDS spread.

Those findings of chapter 2 has important policy implications. Hart and Zingales (2011) suggest using CDS spreads on the long-term debt of large financial institutions to regulate them. The purpose is to contain systemic risk caused by the failure of those too-big-to-fail financial institutions. The idea is that a high CDS spread signals a high default risk of debt issued by the institution. In this case, the financial institution should be required to issue more equity or the regulator should intervene. There may be a tendency to extend this idea to public debt management. That is, one may suggest using sovereign CDS spread to monitor the sovereign credit risk and guide actions to prevent or resolve a sovereign debt crisis. However, a precondition of such a policy proposal is that sovereign CDS spreads should be reliable indicators of the sovereign credit risk. We find that during the run-up to a sovereign debt crisis, the information content of sovereign CDS spreads can be highly distorted by investors' animal spirits. Therefore, it is not advisable to implement public debt management policies based on sovereign CDS spreads.

[1] See Akerlof and Shiller (2009) for a brief review of the history of the interaction between the animal spirits concept and economic theories.

[2] See Farmer (2008) for a survey of ration expectations models of animal spirits.

Chapter 3 of the thesis studies implications of a central bank's monetary policy rules on long-run financial stability. The pioneering work of Kydland and Prescott (1977) and Barro and Gordon (1983) shows that discretionary monetary policy can lead to excessively high inflation rate. One response to this theoretical result is to shift from discretionary monetary policy to monetary policy rules. The best-known monetary policy rule is the Taylor (1993) rule. Taylor (1993) argues that the monetary policy of the United States Federal Reserve under Greespan's chairmanship can be well described by a reaction function of the nominal federal funds rate to inflation and the output gap, the deviation of the logarithm of real gross domestic product from its trend. This reaction function described by Taylor is called the Taylor rule. According to the Taylor rule, other things being equal, the nominal federal funds rate increases by 1.5 percent if the annual inflation rate increases by 1 percent. It increases by 0.5 percent if the output gap increases by 1 percent, other things being equal.

Later theoretical developments link the coefficients of the inflation rate and the output gap in the central bank's reaction function to the stability of the economy. The famous Taylor principle says that if the central bank's policy interest rate adjusts more than one-for-one with inflation, the economy will be stabilized.[3] Bernanke and Gertler (2000) and Bernanke and Gertler (2001) further argue that a central bank's interest rate rule which requires the policy interest rate to react aggressively on inflation is enough to stabilize the asset market as well. This argument is based on financial accelerator models (Kiyotaki and Moore, 1997; Bernanke et al., 1999). In such models, decline in asset prices increases the external finance premium faced by firms. More expensive external borrowing reduces investment and aggregate demand. The decline in aggregate demand in turn reduces inflation. Therefore, asset prices and the general price level always go in the same direction. Asset prices are stabilized if the general price level is stabilized. When both asset prices and the general price level are stabilized, real economy is also stabilized.

An important implication of financial accelerator models is that the central bank is able to handle both the stability of the real economy and financial stability at the same time with only one instrument, the nominal interest rate. In Chapter 3 of this thesis, this

[3] The validity of this statement is a major research topic in macroeconomics. See Clarida et al. (2000), Woodford (2003), Lubik and Schorfheide (2004), Davig and Leeper (2007), Sims and Zha (2006), Farmer et al. (2009), Farmer et al. (2010).

ideal result disappears. In this chapter, there are two different types of investors who have to borrow from financial intermediaries to make their investment. One type invests in the production sector while the other engages in a gambling activity. The central bank's monetary policy affects the loan portfolio faced by financial intermediaries by affecting expected cash flows of both types of borrowers. By lowing the interest rate, the central bank makes debt repayment easier for investors in both production and gambling activities. This encourages entry of both types of investors. More investment can boost the economy in the short run. However, more investors entering the real sector during the boom makes competition tougher and deters future entry. By contrast, because the outcome of the gamble relies on luck rather than previous entry, future entry in the gambling market remains relatively stable. Therefore, in the long run the proportion of gamblers in the pool of loan applicants increases, threatening financial stability. An important policy implication of the model in Chapter 3 is that it is difficult to achieve both the stability of the real economy and financial stability with only one policy instrument, the nominal interest rate. Central banks need to be equipped more than one tools. This justifies the ongoing efforts to strengthen the macro-prudential supervision role of central banks (Orphanides, 2011). Future research should be carried out to find the best policy mix to achieve both the traditional policy goal of price stability and the policy goal of macro-prudential supervision.

While Chapter 3 focuses on monetary policy transmission in a closed economy context, Chapter 4 asks the question whether openness has affected monetary policy transmission. More specifically, we ask whether openness has affected the slope of the Phillips curve. Phillips (1958) finds a negative relationship between the unemployment rate and the growth rate of nominal wage in Britain since 1861. This negative relationship is translated into a tradeoff between inflation and unemployment faced by monetary policy makers. The natural rate hypothesis which appears later says that there is a natural rate of unemployment which is determined by the microeconomic structure. There is a short-run tradeoff between inflation and unemployment. However, the actual unemployment rate cannot deviate from the natural rate in the long run. The output level corresponds to the natural rate of unemployment is called potential output. When the unemployment rate is lower than the natural rate, the output level is higher than the potential output. Therefore, the tradeoff between inflation and unemployment in the Phillips curve is usually expressed as a tradeoff between inflation and the output gap. Modern theoretical models of the Phillips curve are based on microeconomic foundations and address the role of rational expectations.[4] An important implication from those model is that inflation expectations have to be anchored to fight inflation.

In the past few years, there is a debate on whether globalization has contributed to a lower long-run inflation rate. A key channel through which globalization may affect the long-run inflation rate identified by the literature is that it affects the slope of the Phillips curve. Romer (1993), Lane (1997) and Rogoff (2003) argue that more trade openness makes the Phillips curve steeper. In other words, increase in the inflation rate for a given level of output expansion relative to the potential output is larger in a more open economy. Central banks prefer a lower inflation rate and want to keep the output at its potential level. Under the assumption that central banks attach fixed weights to inflation and the output gap in their objective functions, a steeper Phillips curve will make it less attractive for a central bank to fight recession by inflationary policies. A lower inflation bias of the central bank reduces the long-run inflation rate. Models with a benevolent central bank which maximizes the welfare of a representative household can give completely different results. Razin and Loungani (2005) argue that globalization de-links domestic consumption from domestic production. As a result, distortions associated with fluctuations in domestic output gap are reduced. At the same time, domestic production will have less impact on the Consumer Price Index. Therefore, globalization both flattens the Phillips curve and reduces the relative weight of the output gap in central banks' objective functions. Reduction in the long-run inflation rate is due to the reduction in the relative weight of the output gap in central banks' objective functions.

A critique to those models is that the predicted impact of trade openness on the slope of the Phillips curve is not observed in the OECD countries. Chapter 4 of this thesis shows that empirical evidence supporting this critique is based on a wrong parameter homogeneity assumption. More specifically, empirical studies which find trade openness has no effect on the slope of the Phillips curve in the OECD assume that openness has the same effect on the slope of the Phillips curve across countries. Chapter 4 presents theory and evidence that this assumption is wrong. When cross-country heterogeneity is properly modeled, openness is found to have significant effects on the slope of the

[4]See Woodford (2003) for a survey.

Phillips curve in several major OECD countries. This result suggests that it is too early to conclude that globalization has not contributed to the reduction in the long-run inflation rate in the OECD. Woodford (2007) argues that globalization has no impact on monetary control. This claim is too strong. If the tradeoff between inflation and the output gap faced by central banks is changed by globalization, their policy will adjust accordingly. At this stage, it is still not clear what is the optimal monetary policy or policy mix in a global context. A lot more research has to be carried out before giving serious policy suggestions. An interesting direction is to study what are the implications of globalization on the interactions between monetary policy and financial stability.

To summarize, this thesis focuses on three interlinked topics. Chapter 2 studies the determinants of sovereign CDS spreads in Greece, Ireland, Italy, Portugal and Spain during the recent global financial crisis and European debt crisis. The rational expectations assumption of the standard macroeconomic model is relaxed to explain the findings. Chapter 3 introduces a model on the interactions between monetary policy rules and long-run financial stability. In this chapter, we relax the representative agents assumption of the standard macroeconomic model. Chapter 4 reports that globalization has significantly affected the output gap-inflation tradeoff faced by central banks. This result challenges the statement that globalization has no impact on monetary control. I invite interested readers to read each of the chapters for technical details and other interesting findings.

Chapter 2

ANIMAL SPIRITS IN THE EURO AREA SOVEREIGN CREDIT DEFAULT SWAP MARKET[5]

2.1. Introduction

During the European sovereign debt crisis, sovereign credit default swap (CDS) spreads of the Euro countries drew a lot of public attention. The reason is that a country's CDS spread is usually taken as an indicator of that country's sovereign credit risk.[6] In this chapter, we test the reliability of the sovereign CDS spread as an indicator of the sovereign credit risk. More specifically, we test whether changes in variables related to the sovereign credit risk are significant determinants for changes in the sovereign CDS spreads of five Euro-area countries (Greece, Ireland, Italy, Portugal and Spain) in the post-Lehman-Brothers period (from September 15, 2008 to December 19, 2011).

There are a number of empirical studies on the determinants of sovereign CDS spreads in developed countries. Longstaff et al. (2011) find that global financial market conditions significantly affect sovereign CDS spreads of 26 countries, including both developing and developed countries such as Japan and Korea. Dieckmann and Plank (2011) extend their analysis to Western European countries and find that global financial factors also play significant roles there. Moreover, they report that changes in the performance of the financial industry affect changes in the CDS spreads of Western European sovereigns. This finding is consistent with a private-to-public risk transfer hypothesis: prospective government debt necessary to help the distressed financial industry may increase a country's sovereign credit risk. Fontana and Scheicher (2010) focus on Euro area countries

[5]This chapter is coauthored with Hans Blommestein and Sylvester Eijffinger.

[6]In this chapter, we define the sovereign credit risk by the default probability of the sovereign bonds and the associated recovery rate after default.

and also find changes in sovereign CDS spreads are related to global factors. They find that measures of investors' changing risk appetite play a prominent role in the sovereign CDS pricing.

All those previous empirical studies share two features. First, the empirical models are linear. More specifically, there is no regime switching in the models. Second, the covariates are assumed to be exogenous. In this chapter, we show that those two features can bias the statistical inference.

Regime switching can arise from three different theories. Two of those theories are related to different concepts of "animal spirits". In the rational expectations framework, the animal spirits (henceforth we shall call this concept of animal spirits "animal spirits 1") are interpreted as sunspot shocks[7] to investors' expectations (Farmer, 2008). Those sunspot shocks cause multiple equilibria. The economy will be in a good equilibrium if people believe so while the economy will be in a bad equilibrium if people believe it to be bad. Such sunspot-driven multiple equilibria have been used to explain different economic phenomena. They are used to explain excessive volatility in macroeconomic variables such as output and inflation (Clarida et al., 2000; Lubik and Schorfheide, 2004; Davig and Leeper, 2007; Farmer et al., 2010). Diamond and Dybvig (1983) use them to explain bank runs. In international finance, they are used to explain self-fulfilling currency crises (Burnside et al., 2008; Jeanne, 2000). Jeanne and Masson (2000) propose an empirical test for the existence of rational sunspot equilibria in the currency crises context. They prove that the effects of the sunspot shocks are absorbed by discrete jumps in the intercept of a regression of the currency devaluation probability on fundamental variables. Therefore, a test for Markov regime switches in the intercept can be taken as a test for the existence of sunspot equilibria. We argue in Section 4.2 that this test can be applied to the sovereign CDS market *under the rational expectations assumption.*

While the theory of animal spirits 1 predicts regime switches in the intercept of the regression model, an alternative theoretical model under the rational expectations assumption predicts regime switches in the slopes of the regression model. Assuming that investors are rational and there is no sunspot equilibrium, the slopes change if governments change their preferences over different policy objectives. For example, when the

[7] The sunspot shocks are defined as psychological changes which are not related to economic fundamentals.

financial crisis deepens, the weight attached to financial stability may become larger relative to economic growth in governments' objective functions. Anticipating this, rational investors will change their pricing behavior accordingly. Section 4.2 shows that this can lead to regime-dependent slope changes in the CDS spread determination equation.

Under the rational expectations assumption, investors are cognitively unlimited. Therefore, changes in market-based indicators of fundamentals always provide reliable information on the development of fundamental variables. Moreover, the information will be correctly incorporated into sovereign CDS spreads. Those results no longer hold if investors are not cognitively limited. When uncertainties overwhelm the market and there are time constraints for decision-making, the investors rely more on beliefs that are not necessarily based on rational calculations. We call those movements in beliefs of the cognitively limited investors "animal spirits 2" since they are different from the sunspot shocks ("animal spirits 1") in the rational expectations framework. The "animal spirits 2" concept is close to the definition of animal spirits in two recent theoretical papers by De Grauwe (2011a, 2012). In those two papers, agents are not fully rational, that is, they are cognitively limited, and use heuristics rather than rational calculations to make decisions. Agents' sentiments are self-fulfilling because they switch from an optimistic forecast rule to a pessimistic forecast rule if more other agents adopt the pessimistic rule. The widespread pessimistic psychology dampens aggregate demand and eventually leads to a bad outcome. De Grauwe (2011a, 2012) formalizes the concept of "confidence multiplier" of Akerlof and Shiller (2009). According to Akerlof and Shiller (2009), confidence is the belief of cognitively limited agents rather than a sunspot shock to the expectations of perfectly rational agents. De Grauwe (2011a, 2012) does not consider the possibility that agents can change their focus variables in their decision rules if market condition changes. Our definition of "animal spirits 2" allows this possibility.[8] Particularly, when market conditions become very uncertain, agents may drop decision rules based on observable fundamental variables. It is important to point out that ignoring the fundamentals does not mean that investors are irrational. It may be a *boundedly*[9] rational choice by the cognitively limited and imperfectly informed investors. It is because movements in the

[8]Branch and Evans (2007) introduce a model in which agents can change their focus variable in their decision rules. Those agents are taken as econometricians. As pointed out by De Grauwe (2011a), such agents may have better cognitive skills than agents in the real world.

[9]We use the word "bounded" to suggest that agents are cognitively limited whereas perfect rationality means that agents are cognitively unlimited.

observable fundamental variables are driven by market participants whose cognitive abilities are also limited. The information content of those fundamentals is more seriously distorted by the cognitive biases in a more uncertain market. Therefore, observable fundamental variables which are useful when market uncertainty is low can become useless if market uncertainty becomes very high. The theory of animal spirits 2 is consistent with a two-state regime switching model. Under the regime with low market uncertainty, market-based indicators of fundamental variables have significant explanatory power for changes in sovereign CDS spreads. Under the regime with high market uncertainty, all market-based indicators of fundamentals are insignificant.

Above theoretical possibilities for regime switching motivate a form test for regime switching in regression models. Using the quasi-likelihood ratio test developed by Cho and White (2007), we show that the model linearity assumption in previous studies are not valid.

Previous empirical studies on the determination of the sovereign CDS spreads assume that the covariates are exogenous. This assumption rules out the possibility that dynamics in the sovereign CDS spreads may affect fundamental variables. Ruling out such a possibility can be a source of bias. Particularly, it is possible that changes in the CDS spreads will feedback to governments' borrowing costs and affect domestic economic fundamentals.[10] Using a two-step estimation technique developed by Kim (2009), we estimate our regime switching model with instrumental variables and formally test for endogeneity based on the estimation results. Our test suggests that the domestic fundamentals are indeed endogenous in four sample countries (Ireland, Italy, Portugal and Spain). Therefore, compared to the previous studies using ordinary least squares (OLS), our results are more reliable; not only because we model the omitted nonlinearities caused by regime switches but also because we correct for reverse causality.

We find that there is no regime switch in the intercept of the regression equations. This indicates a failure of the joint hypothesis of rational expectations and sunspot equilibria. Therefore, the theory of animal spirits 1 is rejected. There are regime switches in the slopes of the regression equations. A possible explanation is that there is a unique rational expectations equilibrium in which the policy focus of the government changes. The difficulty with this explanation is that there is one regime under which we find that

[10]See OECD (2012).

the sovereign CDS spreads are white noise. That is, they are completely disconnected from all the fundamental variables. The results are better explained by the theory of animal spirits 2. Observable indicators of fundamentals have little value to investors in the sovereign CDS market due to distortions caused by cognitive biases when market uncertainty is high. They are more valuable and used by investors to price the sovereign CDS contracts when market uncertainty is low.

The rest of the chapter is organized as follows: Section 4.2 elaborates on three empirical hypotheses. Section 2.3 introduces the explanatory variables and describes the data. Section 2.4 provides estimates of OLS regression models for the determination of sovereign CDS spreads and tests for regime switching in the models. Section 2.5 shows estimated regime switching models with instrumental variables and tests for endogeneity. Section 2.6 concludes.

2.2. Empirical hypotheses

In the section, we elaborate on three alternative empirical hypothesis for the Euro-area sovereign CDS market.

Hypothesis 1 (animal spirits 1): agents are fully rational and there exist multiple sunspot equilibria.

According to Reinhart and Rogoff (2009), a country's default decision is the result of a cost-benefit analysis. Many countries default on their debts long before they run out of financial resources. Under the rational expectations assumption, Jeanne and Masson (2000) model a country's probability of currency devaluation as a result of its cost-benefit analysis. Due to the similarity, we can apply that model to our sovereign CDS context. More specifically, let us assume that the net benefit function of the government is $B(f_t, d_t)$, where f_t is an index of economic fundamentals, $d_t \equiv \int_0^1 d_t(i)di$ is the average estimate of the probability of default formed by a continuum of investors $i \in [0, 1]$.[11] The net benefit function is increasing in f_t, reflecting the idea that the better the fundamentals are, the higher will be the chance that the government will honor its debt. It is

[11]The two-step estimation approach of Kim (2009) is designed for time series not for panel data analysis. Therefore, we estimate empirical models separately for each country. That is why we only have the time subscript for variables.

decreasing in d_t, suggesting that it is more costly to honor the debt if the investors have higher estimates for the default probability. More specifically, a higher expected default probability increases the interest rate for sovereign borrowing and induces the investors to divert their investment to safer assets, making rollover more difficult (De Grauwe, 2011b).

Investor i expects that the government will default if the net benefit of honoring its debt becomes negative. Therefore, $d_t(i) = Prob[B(f_{t+1}, d_{t+1}) < 0|f_t]$, where $Prob$ denotes probability. Following Jeanne and Masson (2000), we assume that the investors share common knowledge so that we can drop index i in the formula. Under some additional technical assumptions[12], there is a critical value of the fundamental index below which the government will default, given the market estimate of the default probability. Therefore, we can write the average estimate of the default probability as $d_t = Prob[f_{t+1} < f^{*e}|f_t] \equiv F(f_t, f^{*e})$, where f^* is the critical value defined by the equation $B(f^*, d_t) = 0$, the superscript e denotes expectation. Note that f^* is an implicit function of d_t and d_t is a function of f^{*e} , so f^* is a function of f^{*e}, which we denote by $g(f^{*e})$. Under the rational expectations assumption, $f^* = f^{*e}$, so $f^* = g(f^*)$. That is, f^* is a fixed point of the function g. Jeanne and Masson (2000) show that there can be more than one fixed point of g. Their proposition 1 further establishes that if there is more than one fixed point of g, there will be multiple sunspot equilibria. More specifically, there will be n states under which the threshold fundamental index value (denoted by f_s^*, where s is the state index) differs. The probability of default depends not only on the fundamental variables, but also on the transition probabilities from the current to the future states:

$$d_t = \Sigma_{s=1}^{n} q(s_t, s) F(f_t, f_s^*), \tag{2.1}$$

where $q(s_t, s)$ is the transition probability from the current state to state s in the next period, $F(f_t, f_s^*) \equiv Prob[f_{t+1} < f_s^*|f_t]$, where f_s^* is the critical value of the fundamental index under state s.

Following Jeanne and Masson (2000), we assume that the fundamental index is a linear function of the macroeconomic variables relevant for the policy maker's decision.

[12] see Jeanne and Masson (2000) for details.

More specifically, $f_t = \alpha' m_t$, where m_t is a vector of economic fundamentals, α is a vector of constant coefficients and $'$ is a transpose operator. Under this assumption, Jeanne and Masson (2000) show that equation (2.1) can be linearized to the following form:

$$d_t = \delta_{s_t} + \varphi' m_t, s_t = 1, ..., n, \tag{2.2}$$

where δ_{s_t} is a coefficient changing with the state, and φ is a vector of constant coefficients.

Under the rational expectation assumption, the sovereign CDS spread is determined by the default probability of the underlying bond (d_t) and other variables, such as the recovery rate of the defaulted bond and the investors' risk appetite. We write the linearized pricing equation for the sovereign CDS as follows:

$$CDS_t = l + \phi d_t + \chi' \mu_t, \tag{2.3}$$

where CDS_t is the sovereign CDS spread, l and ϕ are constants, χ is a vector of constant coefficients, and μ_t is a vector of determinants for the sovereign CDS spread other than the default probability. Substitute for d_t using equation (2.2), we get

$$CDS_t = \vartheta_{s_t} + \zeta m_t + \chi' \mu_t, \tag{2.4}$$

where $\vartheta_{s_t} \equiv l + \phi\delta_{s_t}$, $\zeta \equiv \phi\varphi'$. Equation (2.4) suggests that under Hypothesis 1, the sovereign CDS spread determination model can be approximated by a Markov regime switching model in which the intercept changes across states but the slopes are always constant.

Hypothesis 2 (changing policy focus): there is a unique *rational* expectations equilibrium, given a particular set of focus fundamental variables of the government. But the focus may change in the sample period.

Hypothesis 2 means that there is only one fixed point for the function g. In this case, $d_t = F(f_t, f^*)$. Its linearized version can be written as

$$d_t = a + b f_t, \tag{2.5}$$

where a and b are constants.[13] If the fundamental index f_t is a linear function of relevant fundamental variables, we will get a constant coefficient CDS pricing model. However, under hypothesis 2, f_t is not a linear function. Instead,

$$f_t = \alpha'_{j_t} m_t, j_t = 1, ..., J, \tag{2.6}$$

where α_{j_t} is a vector of coefficients which change with a discrete state variable j_t, J is the number of possible combinations of target fundamental variables in the objective function of the government, and m_t is the collection of all the potentially relevant variables. Equation (2.6) captures the idea that the set of fundamental variables in the government objective function can change during a turbulent period. More specifically, there is an unobservable latent state variable j whose value governs the changes in the preference of the government. Note that equation (2.6) not only allows changes in the relative weights of the same set of fundamental variables but also allows the set of relevant fundamental variables to change across states. Changes in the set of relevant fundamental variables can be modeled by setting different elements of α_{j_t} to zero under different states. Combining equations (2.3), (2.5), and (2.6), we get

$$CDS_t = \iota + \kappa_{s_t} m_t + \chi' \mu_t, \tag{2.7}$$

where $\iota \equiv l + a\phi$ and $\kappa_{s_t} \equiv b\phi\alpha'_{j_t}$. Thus, under Hypothesis 2, it is the slope vector rather than the intercept that changes across different states. Note that it is possible that not only κ but also some elements of χ change with the state variable.[14] For example, the recovery rate also depends on the cost-benefit analysis of the defaulting government (Reinhart and Rogoff, 2009). Therefore, our reasoning for regime-dependent parameter changes in κ should also be applicable to the coefficients of determinants for the recovery rate.

Hypothesis 3 (animal spirits 2): agents are only boundedly rational. They rely on beliefs which are not related to the observable fundamentals if market uncertainty is high.

The derivation of equations (2.4) and (2.7) depends on the rational expectation as-

[13] Following Jeanne and Masson (2000), we assume that $f_t = \bar{f} + cf_t$, where $\bar{f}$ and c are constants and cf_t is of the first order. Under this assumption, $a = F(\bar{f}, f^*)$ and $b = F_1(\bar{f}, f^*)$.

[14] In this case, we should add a subscript s to χ.

sumption. More specifically, perfect rationality plays three important roles. First, it makes the information content of observable fundamental variables reliable to be used for forecasting the default probability of the sovereign bonds. Second, the forecast of the default probability will be unbiased because the investors are perfectly rational. Third, perfect rationality assures that the CDS spread will correctly incorporate all information on the unbiased forecast of the default probability. If agents are not perfectly rational, those three results will no longer be valid. If market uncertainty is low and cognitive biases are small, a CDS spread determination equation based on those three results may still be a good approximation of reality. In this case, the observable fundamental variables will have explanatory power for the dynamics in the sovereign CDS spreads. However, if market uncertainty is high and cognitive biases are large, it is no longer guaranteed that the observable fundamentals will have explanatory power for the dynamics in the sovereign CDS spreads. It is because the information content of the fundamentals can be highly distorted in a very uncertain environment, and there is no reason to use the incorrect information to price the CDS contract. Therefore, Hypothesis 3 is consistent with a two-state regime switching model.[15] In the less uncertain state, fundamental variables have nonzero coefficients in the CDS spread determination equation. In the more uncertain state, the coefficients of the fundamental variables are zero.

2.3. Variable and data description

2.3.1. The dependent variable: The sovereign CDS spread

The dependent variable in our empirical analysis is the sovereign CDS spread. A CDS contract can be taken as an insurance contract against the credit event specified in the contract.[16] Its spread, expressed in basis points, is the insurance premium the protection buyer has to pay. For example, a CDS spread of 20 basis points means the buyer of credit

[15] In our empirical models, we restrict the number of states to two to save degrees of freedom. Consider only parameters to estimate in the transition matrix. Increasing the number of states from two to three will increase the number of parameters to estimate from 12 to 72 in the two-step regime switching model. Since our sample is relatively small, it is better to restrict the number of states to two. Two is also the typical number of states specified in empirical regime switching models. For example, Jeanne and Masson (2000) use a two-state model.

[16] More precisely, it is a quasi-insurance instrument. See Pan and Singleton (2008) and Dieckmann and Plank (2011) for a more detailed description of the sovereign CDS contract.

protection has to pay the seller an annual amount equal to 0.2 percent of the notional value of the reference debt obligation.[17] There are different credit events against which a sovereign CDS contract can insure. Following Dieckmann and Plank (2011), we consider only the CDS contracts on the credit event "complete restructuring", since it is the standard credit event in the European sovereign CDS contract. The contract maturity we consider is 10 years because the 10-year contract is the most liquid one for the European market. The spreads are quoted in US dollars, the standard currency for European sovereign CDS contracts. Our sample covers weekly data on 10-year government bond CDS spreads from September 15, 2008 to December 19, 2011. Importantly, our sample covers the period after April 2010, which is not covered in the previous studies surveyed in the introduction. Since sovereign debt problems in the sample countries become even more concerned by the public in this period, this extension is particularly interesting.[18] We start the sample from the collapse of Lehman Brothers since the study by Dieckmann and Plank (2011) suggests that European samples before and after the collapse of Lehman Brothers are very different. We include five Euro-area countries (Greece, Ireland, Italy, Portugal and Spain) into our sample. Those five countries are widely believed to have experienced a debt crisis in our sample period. Therefore, it is interesting to ask how reliable are sovereign CDS spreads of those countries as indicators for their sovereign credit risk during the crisis.

2.3.2. The covariates

Table 2.1 summarizes the covariates we use in the regression analysis. As we discussed, the probability of a government's default on its debt depends on the costs and benefits of honoring its debt. Thus, rational investors will use variables that can affect the government's cost-benefit analysis to conjecture the probability of a government default. In addition, they will use this probability of default to price the sovereign CDS contract, an insurance for the sovereign credit risk. Hence, we include variables that are commonly perceived to affect the country's willingness to pay its debt as covariates in the regression analysis. Note that we do not impose the rational expectations assumption for the regression. Rather, we take statistical insignificance of the variables that should have explanatory power to changes in the sovereign CDS spreads under the rational

[17] In our context, the reference debt is the sovereign bond.

[18] See OECD (2012).

expectations assumption as a failure of the assumption.

Theoretically, the state and volatility of the economy may affect a country's willingness to pay its debt. Fiscal reforms necessary to honor the government's debt obligation can impose additional pressure on the already distressed economy. Therefore, when the domestic economy is weak and unstable, the policy maker will be less willing to implement the reforms. Following the literature, we use the domestic stock market return and volatility to proxy the economic state and volatility, respectively. In the rational expectations framework, one should expect the lower the stock market return or the more volatile the return, the higher the sovereign CDS spread, reflecting the unwillingness of the government to take fiscal reforms in an already weak and unstable economy. While Dieckmann and Plank (2011) use the domestic stock price index return, we use the gross return which also includes dividends. This choice is because changes in dividends also contains information on the performance of firms, which affect the performance of the economy. Another domestic variable we consider is the stock market performance of domestic financial firms, the Dow Jones Total Market(DJTM) Financials index. Dieckmann and Plank (2011) argue that this variable measures the private-to-public risk transfer due to the costs of helping the distressed financial industry. That means we should expect a higher sovereign CDS spread when the DJTM financials index is low.

Longstaff et al. (2011) suggest that changes in the global stock and bond markets can explain a large part of the variation in an individual country's sovereign CDS spread. Empirical studies on the European sovereign CDS market (Fontana and Scheicher, 2010; Dieckmann and Plank, 2011) find the same result. For this reason, we also include indicators of developments in the global stock and bond markets as covariates. More specifically, we follow Dieckmann and Plank (2011) to use the EuroStoxx 50 return and MSCI World Financials index as indicators for global stock market developments. We use 10-year German Bund rate and iTraxx Europe corporate CDS spread as indicators for global bond market developments. Dieckmann and Plank (2011) use corporate bond spreads rather than the iTraxx index to proxy European corporate credit spread. The corporate credit spread is not significant in their time series analysis. By contrast, Fontana and Scheicher (2010) find that the iTraxx index has strong explanatory power in the equation for European sovereign CDS spreads. For this reason, we use the iTraxx Europe index as the proxy for European corporate credit spread.

Theoretically, including global variables into the analysis captures the international spillover effect. The European Monetary Union(EMU)-wide stock market performance, EuroStoxx 50 return, is a proxy for the state of the Euro-area economy. Through trade linkages, the economic conditions in the other member countries can affect the home country's economy. This spillover effect need not to be fully captured by the current domestic stock market return due to the fact that a bad union-wide economic condition may affect the home economy with lags. More importantly, in a monetary union, a sovereign country's probability of default is partly affected by the willingness of the other member countries to bail it out, and the other member countries' willingness to pay will depend on their own economic conditions. In this case, a decline in the union-wide economy, proxied by the EuroStoxx 50 return, will increase the sovereign CDS spread. Similarly, a bad state of the world financial industry may affect the willingness of the international community to help an individual sovereign nation out of its debt problem.[19] Therefore, a decline in the World Financials index may increase the home country's sovereign CDS spread. A higher German Bund rate signals a higher rate of economic growth in Germany. This favorable outcome can in turn help improve the economic conditions of the other EMU countries and increase their willingness to help the member countries which have debt problems. Even if Germany's economic growth does not affect other member countries' economic performance, an improvement in its own economy alone can significantly affect the market expectation of defaults by the Euro-area periphery countries. This spillover effect is because Germany plays a leading role in negotiations on the bailout plans. Thus, we expect that an increase in the German Bund rate may reduce the sovereign CDS spreads of the periphery countries. The European corporate CDS spread index, iTraxx, measures the corporate credit spread in Europe. It contains a proxy for the overall state of the European economy since the recovery rates of defaulted corporate bonds increase as the overall business climate improves (Collin-Dufresne et al., 2001). Because lower recovery rates lead to higher corporate CDS spreads, an increase in the iTraxx index implies a deteriorating macroeconomic condition. In this sense, we expect sovereign CDS spreads to be positively related to the iTraxx index. The iTraxx index also contains a proxy for investors' risk appetite.

[19] We use a worldwide proxy for the performance of the financial sector rather than a Euro-area one because the later is not available.

When investors become more risk averse, they will ask for higher credit spread for both corporate bonds and sovereign bonds. This again suggests a positive relationship between iTraxx and the sovereign CDS spreads.

If changes in the iTraxx index fully capture changes in investors' risk appetite, there is no need to include an additional proxy for the risk appetite into the analysis. Fontana and Scheicher (2010) find that the risk appetite proxy constructed from the Chicago Board Options Exchange Market Volatility Index(VIX) is not significant when the iTraxx index is included in the regression. Nevertheless, we add an additional proxy for investors' risk appetite for robustness. More specifically, we use the difference between the implied and realized volatility of EuroStoxx 50 return as the proxy for the global risk premium. This variable captures the pricing of the volatility risk, and therefore contains information on the investors' risk appetite (Longstaff et al., 2011). The implied volatility is the VSTOXX index directly available from Datastream while the realized volatility is estimated by the Garman and Klass (1980) estimator using a rolling 20-day window.

Finally, we include the nominal Euro-US Dollar exchange rate as a covariate. It is measured by the amount of Euros per 100 US dollars. Thus, a higher value means a depreciation of the Euro against the US dollar. We expect a positive sign of this variable. In other words, a depreciation of the Euro increases the sovereign CDS spread. The exchange rate is taken as a global variable since the exchange rate is determined by the macroeconomic fundamentals of the EMU rather than a single member state.

2.3.3. Orthogonalization

Financial asset returns are highly correlated to each other (see Table 2.2). That means including different asset returns into the regression can cause a multicollineararity problem which affects identification. Therefore, it is better to orthogonalize the variables before using them as covariates in the regression. We follow Dieckmann and Plank (2011) to construct the orthogonalized value of a variable as the sum of the estimated intercept and residuals of a regression of that variable on other covariates correlated to it. More specifically, domestic Financials index returns are regressed on the domestic stock market returns and the World financials index return; the World Financials index return is regressed on the global stock market return. Dieckmann and Plank (2011) do not orthogonalize the domestic stock market returns and the European corporate credit spread. Fontana and Scheicher (2010) suggest that orthogonalizing the domestic stock market returns also helps improve identification. Therefore, we orthogonalize the domestic stock market returns by regressing them on the global stock market return and construct the domestic stock market volatility indicators using the orthogonalized series. Alexander and Kaeck (2008) find that changes in the iTraxx index can be explained by changes in VSTOXX and changes in global stock and bond market conditions. Thus, to facilitate identification, we orthogonalize the change in the iTraxx index by regressing it on the change in the VSTOXX index, the global stock market return, the World Financials index and the 10-year German Bund rate.

2.4. OLS regression analysis

Table 2.4 summarizes the estimation results of the following linear OLS regression model.

$$\Delta CDS_t = \Delta x_t^{'}\beta + \epsilon_t, \tag{2.8}$$

where CDS_t is the sovereign CDS spread, x_t is the vector of covariates listed in Table 2.1, ϵ_t is the i.i.d. error term and Δ is a first difference operator. The OLS regressions assume that ϵ_t is independent of x_t. We follow the previous studies to run the regression with first differenced data.[20] This approach facilitates comparison of the results. Consistent with previous studies, our OLS results suggest that changes in the global bond market conditions have strong explanatory power to changes in sovereign CDS spreads. More specifically, increases in the 10-year German Bund rate significantly reduce the sovereign CDS spreads of Ireland, Italy and Spain; increases in the European corporate credit spreads significantly increase the sovereign CDS spreads of Greece, Ireland, Italy and Spain; better Euro-area economic performance (a higher EuroStoxx 50 return) significantly reduces the sovereign CDS spreads of Italy and Spain. These results are also consistent with the theoretical expectation *under the rational expectations assumption*, as we discussed in the last section. Consistent with the private-to-public risk transfer hypothesis, improvement in local financial firms' performance can reduce the sovereign CDS spread. This reduction effect is statistically significant in Italy and Portugal. Signs

[20]See Table 2.3 for descriptive statistics of first differenced data.

of the estimated coefficients of the World Financials index are positive, which is not only different from the finding of Dieckmann and Plank (2011), but also different from the theoretically expected sign we discussed in the last section. However, due to the econometric deficiency of equation (2.8), both the point estimates and the inference based on it are not reliable. Serial independence test results in Table 2.4 suggest that even if there is just one regime, inference based on standard errors reported in Table 2.4 will be distorted. If the single-regime assumption holds, the serial correlation problem can be corrected by using the serial-correlation robust standard errors for inference. However, if the single-regime assumption fails, even the serial independence test results in Table 2.4 will be unreliable.

Testing for regime switching is quite tricky because there are nuisance parameters that are only identifiable under the alternative hypothesis of two regimes but not under the null hypothesis of one regime. More specifically, a single-regime model can be represented in three different ways. First, it can be taken as a model with two regimes with the same regression coefficients. In this case, the probability associated with each regime is not identifiable. In the other two ways, the single-regime model can be taken as a model with two regimes under which the regression coefficients differ but one of the regime happens with zero probability. In such ways of representation, the regression coefficients of the regime which happen with zero probability are not identifiable. In addition, because probabilities cannot be larger than one, there is a boundary condition imposed in the estimation of the regime-switching model. Due to those facts, the typical likelihood ratio test statistics do not follow the usual χ^2 limiting distribution. Cho and White (2007) propose a quasi-likelihood ratio test for regime switching and tabulated critical values at the 5 percent level. Carter and Steigerwald (2011) point out that critical values reported by Cho and White (2007) are based on 10,000 replications, but fewer than 100,000 replications do not produce stable critical values. They provide 5 percent critical values based on 100,000 replications. Table 2.5 reports the quasi-likelihood ratio test statistics for the null of one regime against an alternative of two regimes. Those values are far larger than the critical values tabulated in Carter and Steigerwald (2011). Therefore, the null hypothesis of a single regime is clearly rejected, and we should not make inference based on the OLS model.

2.5. Regime switching model analysis with instrumental variables

Like the OLS model, the standard regime switching models also assume that the error term is independent of the covariates. However, in our context, this assumption may not be plausible. It is possible that the insurance premium of sovereign borrowing affects the borrowing cost and therefore affect the domestic economy. In this case, the local variables are not exogenous and the standard maximum likelihood estimation of a regime switching model will give us biased results. Kim (2009) proposes a two-step maximum likelihood estimator with instrumental variables to solve this problem. Formally, the model can be written as follows:

$$\Delta CDS_t = \Delta x_t^{'} \beta_{S_{1t}} + e_t, S_{1t} = 1, 2, ..., J_1, \tag{2.9}$$

$$\Delta x_t = Z_t^{'} \gamma_{S_{2t}} + \Sigma_{v,S_{2t}}^{1/2} v_t, S_{2t} = 1, 2, ..., J_2, \tag{2.10}$$

where S_{1t} and S_{2t} are unobservable state variables; $Z_t = I_k \otimes z_t$, I_k is a $k \times k$ identity matrix with k being the dimension of x_t, $\otimes$ denotes the Kronecker product[21], and z_t is a $q \times 1$ vector of instrumental variables; $\Sigma_{v,S_{2t}}$ is a $k \times k$ matrix; J_1 and J_2 denote the number of states; the joint distribution of e_t and v_t is

$$\begin{pmatrix} v_t \\ x_t \end{pmatrix} \sim i.i.d.N \left(\begin{pmatrix} 0 \\ 0 \end{pmatrix}, \begin{pmatrix} I_q & \rho_{S_{1t}} \sigma_{e,S_{1t}} \\ \rho_{S_{1t}}^{'} \sigma_{e,S_{1t}} & \sigma_{e,S_{1t}}^2 \end{pmatrix} \right),$$

$\rho_{S_{1t}}$ is a vector of correlation coefficients, and $\sigma_{e,S_{1t}}$ is the standard deviation of e_t. Equation (2.9) is similar to equation (2.8) but now the parameters in β change with the unobservable state variable S_{1t}. The Lucas critique suggests that a regime shift in the policy process governing equation (2.9) can lead to a regime shift in the dynamics of the CDS spread determinants. Therefore, we allow regime shifts in equation (2.10) as well. The unobservable state variable S_{2t} is correlated to S_{1t} according to the Lucas critique. One way to estimate the system composed of equations (2.9) and (2.10) is to specify

[21] Let a_{ij} be the element on the ith row and the jth column of a $m \times n$ matrix A. A$\otimes$B is defined as

$$\begin{pmatrix} a_{11}B & ... & a_{1n}B \\ \vdots & & \vdots \\ a_{m1}B & ... & a_{mn}B \end{pmatrix}$$

.

the joint process of S_{1t} and S_{2t} and estimate the model by a joint maximum likelihood method. However, as pointed out by Kim (2009), such a joint estimation typically has too many parameters to estimate and suffers from the "curse of dimensionality". Furthermore, S_{2t} will be correlated to but different from S_{1t} if there is no perfect policy credibility and the agents have to learn to respond to the policy. Kim (2009) suggests that a two-step estimation approach which ignores the correlation between the state variables suffers less from the "curse of dimensionality". It has better finite sample performance than the joint maximum likelihood estimation when the correlation between S_{1t} and S_{2t} is not perfect. Moreover, it is more robust when the instrument variables are weak. The two-step approach of Kim (2009) first estimates equation (2.10) as a standard regime switching model. This procedure will give consistent estimates for $\gamma_{S_{2t}}$ and $\Sigma_{v,S_{2t}}$ since there are no endogenous covariates in equation (2.10). The elements of the residual vector $\hat{v}_t$ are used as control variables in the second-step estimation of equation (2.9).[22] Kim (2009) proves that this two-step approach will give us consistent estimates for the parameters in equation (2.9).[23]

To save degrees of freedom, we restrict the number of possible states for both S_{1t} and S_{2t} to two. We instrument the local determinants of the CDS spread ($\Delta sdri_t$, $\Delta svol_t$, and $\Delta fdri_t$) by the second and third lags of those local variables and the lagged dependent variable ΔCDS_{t-2} and ΔCDS_{t-3}. Table 2.6 summarizes our two-step estimation results of equation (2.9). Changes in the global bond market conditions (*gbi* and/or *itraxx*) remain to be significant explanatory variables for changes in country-specific sovereign CDS spreads under at least one regime. Moreover, the estimated signs of *gbi* and *itraxx* are consistent with the theory *under the rational expectations assumption*. More specifically, the 10-year German Bund rate (*gbi*) has a negative sign when significant, suggesting that investors expect a lower sovereign credit risk when Germany has a better economic performance. The iTraxx index has a positive sign when significant. As we discussed above, both a worse business climate in the European countries and a higher degree of risk aversion can lead to a higher iTraxx index. Therefore, both a worse economic state of EU and a higher degree of risk aversion can increase the prices of insurances on the sovereign bonds. Similar to the finding by Fontana and Scheicher

[22] See the appendix of this chapter for a brief description of major steps of the second-step estimation.
[23] The second-step standard errors are biased due to the generated regressor problem. The standard errors in the tables are corrected using the method provided by Kim (2009).

(2010), the other proxy for investors' risk appetite, vp, is not significant when the iTraxx index is included as a regressor. The World Financials index is significantly negative under one regime in Italy. This suggests that there is a private-to-public risk transfer in Italy. Under the specific regime, a worse performance of the global financial sector increases the possibility that foreign countries have to spend money to bail out their own financial firms and hence less willing to help the home country. As a result, the sovereign CDS spread increases. Note that it is the performance of the global rather than local financial industry that matters. This finding suggests that compared to the possibility that the Italian government has to bail out its domestic financial firms, the market is more concerned about whether there will be international financial assistance if Italy is in trouble. Under regime 2, $\Delta fgro_t$ turns insignificant while the proxy for domestic economic performance turns significant in Italy. This suggests that under this regime, investors care more about the Italian economy than contingent government debt for bailing out the financial sector. Note that the signs of the estimated coefficients of the World Financials index are positive in some sample countries in some regimes. However, those coefficients are not statistically significant. Hence, it is better to be interpreted as no effect rather than a positive effect. Our results do not support Hypothesis 1, the rational sunspot-equilibria interpretation of the animal spirits. The slopes rather than intercepts are regime-dependent in the sample countries. Hypothesis 2 is also not plausible, though it allows regime-dependent slope changes. Under one of the two regimes, none of the indicators for economic performance and financial health are significant. In this case, Hypothesis 2 will lead to the conclusion that both economic growth and financial stability become unimportant for the government, which is very unlikely to be true. However, our results are consistent with Hypothesis 3. Increasing market uncertainty can enlarge cognitive biases in the market-based indicators for economic and financial health, making them useless for sovereign CDS pricing. Note that the iTraxx index contains a market-based proxy for risk aversion *under the rational expectations assumption*. It becomes insignificant when all observable fundamentals become insignificant. This does not mean that the degree of risk aversion does not matter. Instead, it means that the market-based measure of investors' risk appetite becomes very imprecise when market uncertainty is high and the rational expectations approximation is far from reality.

2.5.1. Tests for endogeneity and serial independence

Kim (2009) suggests that endogeneity of the explanatory variables can be tested by the standard Wald test using the second-step estimation outputs. More specifically, in the two-step estimation, endogeneity is captured by the first-step regression residuals of the endogenous variables on the instrumental variables. These residuals are used in the second-step regression as control variables to eliminate the endogeneity. Therefore, we can test for endogeneity by testing the statistical significance of the first-step residuals in the second-step regression. Formally, the second-step estimation equation can be written as

$$\Delta CDS_t = \Delta x_t^{'}\beta_{S_{1t}} + \hat{v}_t^{'}\theta_{S_{1t}} + \omega_t, S_{1t} = 1, 2, ..., J_1, \tag{2.11}$$

where $\theta_{S_{1t}}$ is a vector of regime-dependent coefficients, $\hat{v}_t$ is the first-step estimate for v_t, and ω_t is an i.i.d. normal random variable given a specific value of S_{1t}. The variance of ω_t changes across regimes. We denote it by $\sigma_{\omega,S_{1t}}$.[24] No endogeneity means $\theta_1 = \theta_2 = ... = \theta_{J_1} = 0$. Under the null hypothesis of no endogeneity, the asymptotic distribution of the Wald statistics $\hat{\theta}' c\hat{o}v(\hat{\theta})^{-1}\hat{\theta}$ is $\chi^2(h)$, where cov denotes the covariance; $\hat{\theta} = [\hat{\theta}_1' = \hat{\theta}_2' = ... = \hat{\theta}_{J_1}']'$ is the vector of estimated values for $\theta_{S_{1t}}$, $S_{1t} = 1, 2, ..., J_1$; h is the dimension of $\hat{\theta}$. Table 2.7 summarizes the Wald test results. The null hypothesis of variable exogeneity is rejected in all sample countries, except Greece. This verifies the importance of controlling for potential endogeneity.

Since we cannot directly apply the Hamilton (1996) test for autoregression to our regime-switching model with endogenous variables, we test for autoregression by adding the lagged dependent variable, ΔCDS_{t-1}, to the second-step equation and test the statistical significance of the autoregressive term. In order to avoid correlation between higher-order lags of ΔCDS_t and ΔCDS_{t-1}, we exclude them from the original instrument variable set. That is, we only use lags of the local variables as instrument variables. Table 2.8 summarizes the estimated coefficients of ΔCDS_{t-1} and their standard errors. The lagged dependent variable is not significant in any sample country under either regime, which suggests no serial correlation in the original model.

[24]See Kim (2009) for details.

2.5.2. The endogeneity of the performance of global financial sector

In the econometric analysis above, we considered only the potential endogeneity of the local variables. Now we consider the potential endogeneity of a global variable: the change in the performance of the global financial sector, $\Delta fgro_t$. Such endogeneity can arise if financial firms outside the home country are highly involved in the trading of the specific country's sovereign CDS contracts.[25] Taking $fgro$ as an additional endogenous variable, we re-estimate the regime switching model. We use the second and third lags of $\Delta sdri_t, \Delta svol_t, \Delta fdri_t, \Delta fgro_t$ and the lagged dependent variable ΔCDS_{t-2} and ΔCDS_{t-3} to instrument the potentially endogenous variables ($\Delta sdri_t, \Delta svol_t, \Delta fdri_t, \Delta fgro_t$). We test the endogeneity of $fgro$ based on the new estimation results. As we mentioned in the last subsection, the test for endogeneity is equivalent to the test for the statistical significance of the corresponding first-stage residuals. Table 2.9 summarizes the test results. Those results suggest that changes in the Irish and Portuguese sovereign CDS spreads have significantly affected changes in the performance of financial firms outside those two countries at least under one regime. Table 2.10 reports the estimation results for Ireland the Portugal, taking $fgro$ as an endogenous variable. The previous result that changes in the fundamental variables do not explain changes in the Irish or Portuguese sovereign CDS spreads under regime 2 is unchanged. This means that the type-2 animal spirits of investors are indeed the driver of changes in the Irish and Portuguese sovereign CDS spreads under regime 2. Changes in the Euro-Dollar rate and the iTraxx index significantly affect changes in the Irish sovereign CDS spread under regime 1. More specifically, a depreciation of the Euro relative to the US Dollar and an increase in the European corporate CDS spread lead to an increase in the Irish sovereign CDS spread. The significant positive sign of the iTraxx index suggests that either a worse business climate increases the sovereign credit risk or a higher degree of risk aversion increases the insurance premium for the sovereign borrowing. In Portugal, under regime 1, the 10-year German Bund rate appears to be the only significant fundamental driver of the sovereign CDS spread. The negative sign of *gbi* suggests that a larger increase in the German growth rate implies a higher increase in the probability that the EMU will provide financial support to the Portugal government if it is in trouble.

[25] See OECD (2012).

2.6. Conclusion

We have studied the determinants of changes in the sovereign CDS spreads of five Euro-area countries (Greece, Ireland, Italy, Portugal and Spain) after the failure of Lehman Brothers. Two distinct regimes under which the coefficients of the determinants differ are identified.

On the one hand, under regime 2, the usual determinants of changes in the sovereign CDS spreads of Greece, Ireland, and Portugal lose their explanatory power.[26] We argue that the animal spirits; that is, the psychological movements of cognitively limited investors are the key drivers of the sovereign CDS spreads in such situations. This has important implications for both policy makers and academic researchers. As a widely-used indicator for the sovereign credit risk, the sovereign CDS spread can be highly distorted in the sense that it can be completely disconnected from the country's fundamental economic movements. In the rational expectations framework, the existence of non-fundamental determinants of sovereign CDS spreads does not necessarily mean that the CDS spreads cannot predict sovereign defaults. That is because non-fundamental sunspot shocks to investors' expectation can lead to self-fulfilling sovereign debt crises. If the market believes that a debt crisis is under way, it will happen. And if the market participants are perfectly rational, the sovereign default probabilities will be correctly included in the pricing of the corresponding sovereign CDS spreads. However, our empirical results do not support the story of rational self-fulfilling debt crises. Rather, there are periods in which the boundedly rational market participants fail to price the sovereign credit risk correctly. In this case, the sovereign CDS spread is not very useful for evaluating the sovereign credit risk.

On the other hand, our results also suggest that there are periods in which the fundamental variables matter. In addition, the estimated effects of those fundamental variables reflect at least bounded rationality of the market participants. More specifically, we find that in the periods when the investors behave more rationally, the global bond market conditions are particularly important for the pricing of the sovereign CDS spreads. A better economic prospect of Germany or a better European-wide business climate implies a higher chance that other union members will be willing to provide financial support

[26]Under this regime, the market is more turbulent in the sense that the conditional variance of changes in the sovereign CDS spreads is larger than under regime 2. See Tables 2.6 and 2.10.

for the fiscally distressed countries. Therefore, the market confidence on sovereign borrowing will be enhanced and sovereign CDS spreads decrease. A depreciation of the Euro against the US dollar significantly increases Italian and Spanish sovereign CDS spreads. This result also suggests that the Euro-area economic prospect can affect an individual member country's sovereign credit risk. It is because the depreciation of the Euro can signal a weaker Euro-area economy. Another interesting finding is that there are periods in which traders of the Italian sovereign CDS contracts are concerned by the performance of the global financial industry. A worse performance of the global financial industry increases the probability that foreign countries will have to bail out their own financial sectors and become less willing to provide financial aids for Italian fiscal reforms. By contrast, in some other periods, investors in the market of Italian sovereign CDS contracts only cares about the domestic economic performance of Italy, and pay little attention to the performance of the global financial sector. The reason for this change in investors' concern can be an interesting topic for future research.

Table 2.1: Variable definitions

Variable	Definition
forex	Nominal Euro to US Dollar exchange rate, the amount of Euros per 100 US Dollars
stoxx	EuroStoxx 50 return (orthogonalized), percentage point
gbi	10-year benchmark German Bund interest rate, basis point
itraxx	iTraxx Europe 10-year CDS spread (orthogonalized), basis point
vp	Volatility risk premium, percentage point
fgro	MSCI World Financials index return (orthogonalized), percentage point
sdri	DJTM domestic stock market return (orthogonalized), percentage point
svol	GARCH(1,1) Domestic stock market volatility, percentage point
fdri	DJTM Financials index return (orthogonalized), percentage point

Notes: All data are from Datastream.
See the texts for detailed description on the orthogonalization of variables.
The volatility risk premium is proxied by the difference between the implied volatility and the Garman-Klass realized volatility of EuroStoxx 50.

Table 2.2: Correlation between stock market returns

Greece					Ireland				Italy			
	stoxx	fgro	sdri	fdri	stoxx	fgro	sdri	fdri	stoxx	fgro	sdri	fdr
stoxx	1.00				1.00				1.00			
fgro	0.84	1.00			0.84	1.00			0.84	1.00		
sdri	0.68	0.59	1.00		0.77	0.77	1.00		0.95	0.81	1.00	
fdri	0.61	0.53	0.95	1.00	0.59	0.63	0.74	1.00	0.89	0.78	0.96	1.00
Portugal					Spain							
	stoxx	fgro	sdri	fdri	stoxx	fgro	sdri	fdri				
stoxx	1.00				1.00							
fgro	0.84	1.00			0.84	1.00						
sdri	0.78	0.64	1.00		0.92	0.77	1.00					
fdri	0.58	0.51	0.73	1.00	0.88	0.78	0.97	1.00				

Notes: Correlation coefficients are calculated using non-orthogonalized data.

Table 2.3: Descriptive statistics

	CDS	forex	stoxx	gbi	itraxx	vp	fgro	sdri	svol	fdri
Greece										
Mean	42.81	0.05	-0.02	-1.39	0.12	0.00	-0.01	-0.03	-0.31	-0.04
Median	8.00	-0.01	-0.68	-1.5	0.22	0.18	0.03	-0.32	3.87	0.19
Maximum	3648.5	5.18	26.16	35.20	19.67	20.43	14.85	16.50	294.61	22.43
Minimum	-1622.96	-4.20	-13.44	-36.20	-30.00	-22.04	-11.72	-16.09	-393.21	-15.00
Standard deviation	397.60	1.40	5.88	12.59	7.19	5.34	3.83	6.89	47.40	3.88
Ireland										
Mean	3.47	0.05	-0.02	-1.39	0.12	0.00	-0.01	0.06	-4.28	-0.10
Median	2.94	-0.01	-0.68	-1.5	0.22	0.18	0.03	-0.05	-0.67	-0.35
Maximum	291.44	5.18	26.16	35.20	19.67	20.43	14.85	27.36	6.35	74.67
Minimum	-305.68	-4.20	-13.44	-36.20	-30.00	-22.04	-11.72	-26.38	-54.16	-76.52
Standard deviation	51.99	1.40	5.88	12.59	7.19	5.34	3.83	4.93	10.41	17.04
Italy										
Mean	2.66	0.05	-0.02	-1.39	0.12	0.00	-0.01	0.01	0.05	0.00
Median	1.25	-0.01	-0.68	-1.5	0.22	0.18	0.03	-0.02	-2.71	0.25
Maximum	132.22	5.18	26.16	35.20	19.67	20.43	14.85	7.35	156.34	8.48
Minimum	-101.55	-4.20	-13.44	-36.20	-30.00	-22.04	-11.72	-7.43	-19.96	-6.49
Standard deviation	25.85	1.40	5.88	12.59	7.19	5.34	3.83	1.99	16.34	2.64
Portugal										
Mean	5.48	0.05	-0.02	-1.39	0.12	0.00	-0.01	-0.01	-1.46	-0.05
Median	2.69	-0.01	-0.68	-1.5	0.22	0.18	0.03	-0.05	-1.45	0.20
Maximum	354.84	5.18	26.16	35.20	19.67	20.43	14.85	13.81	25.08	22.95
Minimum	-287.39	-4.20	-13.44	-36.20	-30.00	-22.04	-11.72	-10.21	-22.73	-22.36
Standard deviation	56.81	1.40	5.88	12.59	7.19	5.34	3.83	3.38	5.78	5.79
Spain										
Mean	2.04	0.05	-0.02	-1.39	0.12	0.00	-0.01	-0.01	-0.74	0.01
Median	1.50	-0.01	-0.68	-1.5	0.22	0.18	0.03	-0.13	-1.83	-0.01
Maximum	91.06	5.18	26.16	35.20	19.67	20.43	14.85	9.35	41.26	6.33
Minimum	-107.15	-4.20	-13.44	-36.20	-30.00	-22.04	-11.72	-9.95	-16.17	-7.60
Standard deviation	24.52	1.40	5.88	12.59	7.19	5.34	3.83	2.48	7.16	2.06

Notes: All data are first differenced.

Table 2.4: OLS results

	Greece	Ireland	Italy	Portugal	Spain
constant	0.40	0.02	0.02	0.07	0.02
	(0.30)	(0.04)	(0.02)	(0.04)	(0.01)
forex	-2.70	3.99	2.98**	1.23	2.43**
	(26.43)	(3.29)	(1.32)	(3.16)	(1.28)
stoxx	-10.07	-0.35	-0.69**	-1.28	-0.80**
	(7.26)	(0.87)	(0.36)	(0.87)	(0.34)
gbi	-0.46	-0.75**	-0.55***	-0.63	-0.40***
	(2.81)	(0.35)	(0.14)	(0.34)	(0.13)
itraxx	10.01**	1.18**	0.77***	0.55	1.01***
	(4.49)	(0.55)	(0.24)	(0.55)	(0.21)
vp	2.05	1.75	0.60	1.56	0.57
	(8.16)	(0.99)	(0.40)	(0.97)	(0.39)
fgro	4.24	1.43	0.48	2.71***	0.74
	(8.27)	(1.07)	(0.42)	(0.99)	(0.40)
sdri	-8.04	-0.27	-2.41***	-5.76***	-2.45***
	(4.52)	(0.83)	(0.92)	(1.20)	(0.67)
svol	-0.00	0.06	0.18	1.74***	0.53**
	(0.66)	(0.36))	(0.10)	(0.66)	(0.23)
fdri	-4.19	-0.24	-1.64***	-2.12***	-0.85
	(8.06)	(0.23)	(0.61)	(0.63)	(0.76)
Adjusted R-squared	0.29	0.16	0.43	0.32	0.42
Serial independence	0.56	0.00	0.00	0.13	0.03

Notes: Standard errors in parentheses. ***,** denotes significance at one and five percent level, respectively.
Serial independence is the Lagrange Multiplier (LM) test p value for serial correlation up to two orders.

Table 2.5: Tests for regime switching

	Greece	Ireland	Italy	Portugal	Spain
test statistics	462.6	115.1	75.96	157.8	75.48

Notes: Test statistics are Cho and White (2007) Quasi-Likelihood Ratio test statistics.
Null hypothesis: one regime.

Table 2.6: Regime switching model results-local variables instrumented

	Greece		Ireland		Italy		Portugal		Spain	
	1	2	1	2	1	2	1	2	1	2
constant	0.06	2.01	0.01	0.38	-0.01	0.03	0.04	0.22	0.01	0.00
	(0.04)	(3.64)	(0.03)	(6.05)	(0.01)	(0.06)	(0.02)	(1.31)	(0.02)	(0.03)
forex	5.47	-38.25	3.93	3.82	2.47***	2.19	1.63	12.94	2.67**	2.26
	(4.25)	(395.50)	(2.20)	(653.70)	(0.88)	(7.08)	(2.04)	(99.84)	(1.31)	(3.19)
stoxx	-1.23	-58.30	-0.56	(-6.67)	-0.17	-0.05	-0.58	-10.46	0.02	-1.10
	(1.20)	(89.51)	(0.51)	(195.20)	(0.35)	(1.51)	(0.61)	(25.06)	(0.50)	(0.83)
gbi	-0.84**	13.08	-0.34	-2.39	-0.27**	-0.95	-0.62***	(3.27)	-0.28	-0.74**
	(0.41)	(36.99)	(0.23)	(68.41)	(0.13)	(0.53)	(0.19)	(10.20)	(0.25)	(0.34)
itraxx	0.94	30.43	1.23***	-3.70	0.64***	1.79	0.72**	4.51	0.39	1.94***
	(0.79)	(33.88)	(0.32)	(199.50)	(0.23)	(0.96)	(0.36)	(10.29)	(0.27)	(0.47)
vp	-0.30	16.76	0.05	2.14	-0.16	1.67	0.29	2.36	0.10	0.87
	(1.20)	(140.90)	(0.54)	(267.00)	(0.44)	(1.13)	(0.58)	(36.34)	(0.59)	(0.87)
fgro	0.25	6.18	0.18	5.10	-0.97***	1.30	0.84	-1.61	-0.45	1.34
	(1.25)	(104.10)	(0.65)	(192.00)	(0.32)	(2.15)	(0.70)	(28.37)	(0.54)	(1.12)
sdri	0.02	-36.01	-1.89	21.13	1.04	-16.78**	-2.17	2.76	1.06	1.49
	(1.93)	(74.62)	(1.08)	(746.50)	(1.48)	(8.61)	(1.64)	(84.37)	(1.61)	(3.36)
svol	0.24	-0.69	-0.22	101.60	-0.06	1.21	0.65	-11.70	0.15	0.02
	(0.25)	(6.99)	(0.28)	(781.00)	(0.09)	(1.11)	(0.70)	(39.40)	(0.51)	(0.67)
fdri	0.76	-6.64	0.26	-16.33	0.18	2.98	-0.99	-3.00	-0.21	-0.98
	(4.82)	(64.09)	(0.34)	(102.80)	(1.02)	(5.22)	(0.76)	(45.02)	(1.44)	(5.74)
p_{ii}	0.97	0.91	0.97	0.89	0.93	0.94	0.97	0.87	0.98	1.00
σ_ω	0.39	7.00	0.22	0.55	0.08	0.16	0.20	0.49	0.08	0.18

Notes: Standard errors in parentheses. ***,** denotes significance at one and five percent level respectively.
p_{ii} denotes the probability of staying under regime i next period if i is the current regime.

Table 2.7: Endogeneity tests (local variables only)

	Greece	Ireland	Italy	Portugal	Spain
Wald statistics	1.36	20.39	14.25	13.98	16.97
p value	0.97	0.00	0.03	0.03	0.01

Notes: Testing for endogeneity of the local variables, taking the global variables as exogenous.

Table 2.8: Serial correlation tests for the regime switching model

	Greece	Ireland	Italy	Portugal	Spain
regime 1	0.01	-0.03	0.08	0.07	0.09
	(0.05)	(0.09)	(0.10)	(0.06)	(0.12)
regime 2	-0.21	1.40	-0.33	0.35	-0.02
	(0.73)	(27.4)	(0.27)	(0.22)	(0.34)

Notes: Estimated coefficients of ΔCDS_{t-1} with standard errors in parentheses. ***,** denotes significance at one and five percent level, respectively.

Table 2.9: Tests for the endogeneity of $fgro$

	Greece	Ireland	Italy	Portugal	Spain
regime 1	-0.11	-0.01	0.00	0.07**	-0.01
	(0.14)	(0.04)	(0.02)	(0.03)	(0.02)
regime 2	-2.14	1.18**	0.00	0.09	-0.06
	(8.08)	(0.06)	(0.06)	(0.34)	(0.05)

Notes: Standard errors in parentheses. ***,** denotes significance at one and five percent level, respectively.

Table 2.10: Regime switching model results-local variables and $fgro$ instrumented

	Ireland		Portugal	
	regime 1	regime 2	regime 1	regime 2
constant	0.01	0.02	0.02	0.11
	(0.02)	(1.11)	(0.02)	(0.43)
forex	4.38***	-0.86	3.22	-5.26
	(1.68)	(122.8)	(1.79)	(35.73)
stoxx	-0.71	1.40	-0.37	-4.36
	(0.48)	(34.61)	(0.54)	(12.61)
gbi	-0.34	-1.67	-0.49**	-0.91
	(0.18)	(10.85)	(0.20)	(4.02)
itraxx	1.07***	0.74	0.59	4.85
	(0.29)	(21.62)	(0.32)	(6.75)
vp	-0.10	4.51	0.24	3.57
	(0.48)	(27.53)	(0.56)	(13.63)
fgro	-0.87	25.38	-1.14	7.25
	(0.79)	(201.3)	(1.51)	(25.01)
sdri	-0.16	-1.00	0.57	-3.35
	(0.20)	(132.1)	(0.61)	(11.37)
svol	-0.30	5.39	-0.91	-1.37
	(0.21)	(51.48)	(0.96)	(33.96)
fdri	0.16	-40.80	-0.60	1.55
	(0.94)	(188.7)	(1.02)	(39.63)
p_{ii}	0.98	0.95	0.96	0.92
σ_ω	0.20	0.65	0.15	0.55

Notes: Standard errors in parentheses. ***,** denotes significance at one and five percent level respectively.
p_{ii} denotes the probability of staying under regime i in the next period if i is the current regime.

2.7. Appendix

In this appendix, we show the major steps of the second-step estimation for our two-state model. Our purpose is to estimate $\beta_{S_{1t}}$, $\theta_{S_{1t}}$, $\sigma_{e,S_{1t}}$ and p_{ij}, the transition probability from state i to state j. From equation (2.10), we have

$$\hat{v}_t = inv(\hat{\Sigma}^{1/2}_{v,S_{2t}})(\Delta x_t - Z'_t \hat{\gamma}_{S_{2t}}), \tag{2.12}$$

where $inv(\cdot)$ denotes the inverse, and $\hat{\Sigma}^{1/2}_{v,S_{2t}}$ and $\hat{\gamma}_{S_{2t}}$ denote the first-step estimates for $\Sigma^{1/2}_{v,S_{2t}}$ and $\gamma_{S_{2t}}$, respectively.

Using equations (2.11) and (2.12), we can derive the conditional density function of ΔCDS_t for given values of S_{1t} and S_{2t}. More specifically, for $j_1 = 1, 2$ and $j_2 = 1, 2$, the density functions can be represented as: $f(\Delta CDS_t|\Delta Z_t, \Delta x_t, S_{1t} = j_1, S_{2t} = j_2; \lambda_1, \hat{\lambda}_2) = \frac{1}{\sqrt{2\pi\sigma^2_{\omega,j_1}}} exp\{-\frac{1}{2\sigma^2_{\omega,j_1}}\{\Delta CDS_t - x'_t\beta_{j_1} - [inv(\hat{\Sigma}^{1/2}_{v,j_2})(\Delta x_t - Z'_t\hat{\gamma}_{j_2})]'\theta_{j_1}\}^2\}$, where λ_1 denotes the vector of parameters to be estimated in the second step, and $\hat{\lambda}_2$ denotes the vector of estimated parameters in the first step.

Using the standard smoother for the regime switching model, we can get, from the first-step estimation, $Prob(S_{2t} = 1|\Delta\tilde{x}_T)$ and $Prob(S_{2t} = 2|\Delta\tilde{x}_T)$, where $\Delta\tilde{x}_t$ denotes the historical information on Δx until time t, T is the end of the sample period.[27] We can calculate the conditional densities for $j_1 = 1, 2$: $f(\Delta CDS_t|\Delta Z_t, \Delta x_t, S_{1t} = j_1; \lambda_1, \hat{\lambda}_2) = f(\Delta CDS_t|\Delta Z_t, \Delta x_t, S_{1t} = j_1, S_{2t} = 1; \lambda_1, \hat{\lambda}_2) \times Prob(S_{2t} = 1|\Delta\tilde{x}_T) + f(\Delta CDS_t|\Delta Z_t, \Delta x_t, S_{1t} = j_1, S_{2t} = 2; \lambda_1, \hat{\lambda}_2) \times Prob(S_{2t} = 2|\Delta\tilde{x}_T)$.

Denote the historical information on ΔCDS_t until period $t-1$ by $\Delta\widetilde{CDS}_{t-1}$. If $Prob(S_{1t} = j_1|\Delta\widetilde{CDS}_{t-1}, \Delta\tilde{x}_T)$ is known, we can calculate the predictive density of ΔCDS_t by the following equation:

$f(\Delta CDS_t|\Delta\widetilde{CDS}_{t-1}, \Delta x_t; \lambda_1, \hat{\lambda}_2) = f(\Delta CDS_t|\Delta Z_t, \Delta x_t, S_{1t} = 1; \lambda_1, \hat{\lambda}_2) \times Prob(S_{1t} = 1|\Delta\widetilde{CDS}_{t-1}, \Delta\tilde{x}_T) + f(\Delta CDS_t|\Delta Z_t, \Delta x_t, S_{1t} = 2; \lambda_1, \hat{\lambda}_2) \times Prob(S_{1t} = 2|\Delta\widetilde{CDS}_{t-1}, \Delta\tilde{x}_T)$.[28]

However, we do not know $Prob(S_{1t} = j_1|\Delta\widetilde{CDS}_{t-1}, \Delta\tilde{x}_T)$. Given initial values $Prob(S_{10} = j_1|\Delta\widetilde{CDS}_0, \Delta\tilde{x}_T)$, we can calculate the filtered probabilities as follows: $Prob(S_{1t} = 1|\Delta C\tilde{D}S_{t-1}, \Delta\tilde{x}_T) = p_{11}Prob(S_{1,t-1} = 1|\Delta\widetilde{CDS}_{t-1}, \Delta\tilde{x}_T) + p_{21}Prob(S_{1,t-1} = 2|\Delta\widetilde{CDS}_{t-1}, \Delta\tilde{x}_T)$.

[27] See Hamilton (1994) for details on the standard regime switching model.
[28] Note that in our model, Z_t includes past values of CDS_t and x_t.

Similarly, $Prob(S_{1t} = 2|\Delta C\tilde{D}S_{t-1}, \Delta\tilde{x}_T) = p_{12}Prob(S_{1,t-1} = 1|\Delta\widetilde{CDS}_{t-1}, \Delta\tilde{x}_T) + p_{22}Prob(S_{1,t-1} = 2|\Delta\widetilde{CDS}_{t-1}, \Delta\tilde{x}_T)$.

The probabilities can be updated using the following equation:

$Prob(S_{1t} = j_1|\Delta\widetilde{CDS}_t, \Delta\tilde{x}_T) = \frac{f(\Delta CDS_t|\Delta Z_t, \Delta x_t, S_{1t}=j_1;\lambda_1,\hat{\lambda}_2)\times Prob(S_{1t}=j_1|\Delta\widetilde{CDS}_{t-1},\Delta\tilde{x}_T)}{f(\Delta CDS_t|\Delta\widetilde{CDS}_{t-1},\Delta x_t;\lambda_1,\hat{\lambda}_2)}$, where $j_1 = 1, 2$.

Iterating the procedure listed above, we can get the log likelihood function to be maximized.

Chapter 3

MONETARY POLICY RULES, ADVERSE SELECTION AND LONG-RUN FINANCIAL RISK[29]

3.1. Introduction

Taylor (2009) suggests that government policies could be sources of financial crises. In this chapter, we focus on the impact of one particular type of government policy, that is, central bank's interest rate policy, on financial stability. More specifically, we investigate how the interactions between various shocks and central bank's interest rate rules dynamically affect the adverse selection problem faced by financial intermediaries.

To that end, we build a stochastic dynamic general equilibrium model with two types of borrowers. One is a gambler, who borrows to invest in a fixed-supply gambling asset. The other is an entrepreneur, who borrows to pay the set-up cost for production. Borrowers are protected by the limited liability law. Limited liability together with fixed-supply can generate a bubble in the gambling asset market (Allen and Gale, 2000). When there is a bubble, lending to gamblers generates expected losses. In this case, there are two reasons why the gamblers still get loans from financial intermediaries (Barlevy, 2008). First, lending to entrepreneurs generates expected profits. Second, there is no screening between gamblers and entrepreneurs.[30] Without screening in the financial intermediation sector, a persistent decrease in the proportion of entrepreneurs in the

[29]This chapter is coauthored with Hans Blommestein and Sylvester Eijffinger.

[30]Empirical studies (Giot and Schwienbacher, 2003; Bertoni et al., 2011) suggest that even venture capital firms which are supposed to have a strong ability to select good borrowers do not really select good firms. Reinhart and Rogoff (2009) suggest that the expansion of the financial intermediation sector in the run-up to crises causes overcapacity in the financial industry. Since many new intermediaries enter the market with less experience during the expansionary period, one should expect a weaker average ability to screen the borrowers. Moreover, the theoretical model of Dell'Ariccia and Marquez (2006) shows that financial intermediaries will optimally choose not to screen the borrowers if the number of new loan applicants is sufficiently large.

borrower pool can accumulate huge losses for financial intermediaries. This chapter links the central bank's interest rate policy to changes in the borrower pool in a general equilibrium framework. More specifically, a change in the interest rate affects liabilities of both types of borrowers in the same way, but affects the payoffs of the assets they buy in a different way. On the one hand, the payoff of the gambling asset is exogenously determined by the lotteries. On the other hand, the payoffs of the firms set up by the entrepreneurs are endogenously determined by a number of factors including the interest rate. Therefore, a change in the interest rate can disproportionately change the expected cash flows for gamblers and entrepreneurs. The difference in the changes of expected cash flows leads to a difference in the market entry decisions, which changes the proportion of entrepreneurs in the borrower pool.

The key result of our model is that the central bank's interest rate policy can reduce the riskiness of the loan portfolio in the short run, while persistently increase the riskiness in the long run. More specifically, by lowering the interest rate, the central bank makes debt repayment easier for both entrepreneurs and gamblers. This encourages entry of both types of investors. Our quantitative analysis suggests that entry of entrepreneurs may increase more than entry of gamblers in the short run, which means that the proportion of entrepreneurs in the pool of new loan applicants increases in the short run. Since loans to entrepreneurs are less risky than loans to gamblers, the loan portfolio becomes less risky in the short run. However, more entry of entrepreneurs intensifies competition in the production sector and reduces their future profits. This deters entrepreneurial entry in the long run. By contrast, future payoffs of the lotteries are exogenously determined and are not affected by the current-period entry of gamblers.[31] Therefore, the proportion of entrepreneurs persistently stays at low levels in the long run.

Taylor (2009) argues that deviations from the Taylor rule can be a source of financial crisis. We find that expansionary monetary policy shocks can persistently worsen the borrower pool faced by financial intermediaries in the long run. This is consistent with Taylor's argument. However, quantitative results of our model also suggest that sticking to the Taylor rule is not sufficient to eliminate financial crises. Actually, if the economy

[31] Strictly speaking, not only future payoffs but also asset prices affect entry decisions. Competition can push up the price of the gambling asset. However, it also pushes up the cost of entering the production sector, which increases the value of the firms. Therefore, competition-induced changes in asset prices are limited in relative terms. As a result, the effect of such changes in relative asset prices on the proportion of entrepreneurs is also limited.

is hit by a negative productivity shock, a central bank which deviates from the Taylor rule by not reacting to output fluctuations can reduce long-run financial risk.

The financial accelerator model (Bernanke et al., 1999) also links productivity shocks to financial intermediation in a general equilibrium macroeconomic framework.[32] However, there is no distinctive difference between the short-run and long-run effects of shocks in the financial accelerator model. As we discussed, the distinction between the short run and long run is important. Moreover, the financial accelerator model considers only one type of borrower (the entrepreneur) and assumes that the number of entrepreneurs is fixed. Our model instead features endogenous entry of both gamblers and entrepreneurs. It enables us to study the dynamic changes in the borrower pool faced by financial intermediaries.

Our modeling of the production sector is related to the macroeconomic model with endogenous firm entry by Ghironi and Melitz (2005) and Bilbiie et al. (2008), but differs in several important respects. In Ghironi and Melitz (2005) and Bilbiie et al. (2008), there are no financial frictions and therefore no role for financial intermediation. In our model, there is a financial friction and firms must rely on financial intermediaries to buy goods necessary to start their business. This specification enables us to study the feedback effect from the real sector to the financial sector. In Ghironi and Melitz (2005) and Bilbiie et al. (2008), firms exit exogenously at a constant rate. In our model, the exit of firms is endogenously determined by their ability to repay their debt. One particularly important difference is that nominal wage is sticky in our model, whereas it is flexible in Ghironi and Melitz (2005) and Bilbiie et al. (2008). Bilbiie et al. (2008) find that entrepreneurial entry initially decreases after an expansionary monetary policy shock. This happens because output expansion created by the interest rate shock pushes up the real wage. The higher real wage not only makes entrepreneurial entry more costly but also decreases profits after entry. Therefore, fewer entrepreneurs want to enter. However, as suggested by Rotemberg (2008), if the nominal wage is sticky, the rise in the real wage will be more modest. As a result, profits could rise rather than fall. Therefore, entrepreneurial entry can rise despite the increase in entry cost. Since entrepreneurial entry affects the borrower pool faced by financial intermediaries, it is crucial to model

[32] See Allen and Gale (2007) for a survey of partial equilibrium models which also link real shocks to financial stability.

wage setting in a more realistic way.

A popular claim in media and policy discussions is that speculation in the secondary financial market is a source of financial crisis. Very often this claim is used to justify financial suppression, for example through restrictions on secondary market trading. While fire sales in the secondary market can trigger a financial crisis (Allen and Gale, 2004, 2007), banning secondary market trading is far from a justified solution to prevent financial crises. Our model suggests that productivity shocks originating from the real sector can lead to a financial crisis even if there is a ban on secondary market trading of the gambling asset.

We proceed as follows. Section 2 introduces the model. Section 3 solves the model. Section 4 displays and discusses impulse responses of the model under different shocks. Section 5 studies the robustness of the qualitative results to sticky interest rate pass-through. Section 6 concludes.

3.2. The model

To facilitate the comparison of impulse responses in our model to those in the standard literature, we incorporate nominal (goods price stickiness and wage stickiness) and real frictions (habit formation in consumption and monopolistic competition in production) in the standard new Keynesian model. Aggregation is very difficult if we have both price stickiness and endogenous entry and exit of firms in one sector. Therefore, we separate those two features into two different sectors. First, we have a consumption goods sector with sticky prices and a fixed mass of firms. Second, we have an intermediate goods sector with flexible prices and endogenous entering and exiting of firms. The consumption goods producers use intermediate goods for production while the intermediate goods producers use labor for production. Entering entrepreneurs in the intermediate goods sector must hire labor and buy consumption goods to set up their firms. The entry cost in terms of wage payment is covered by the shareholders or households. The entry cost in terms of consumption goods is covered by loans from financial intermediaries. Besides entrepreneurs, financial intermediaries also face another type of borrower, the gambler. Gamblers use the borrowed amount to buy an asset of which the payoff is completely exogenously determined by lotteries. In each period, financial intermediaries receive an installment repayment from the borrowers if there is no default. Households make decisions on labor supply, goods consumption, investment in deposits and new stocks of firms in the intermediate goods sector. They receive profits and wage payments from the firms and interest payments from the financial intermediaries. Finally, there is a central bank which sets nominal money market interest rates. Figure 3.1 summarizes the interrelationships between the agents.

3.2.1. Firms

Consumption goods producers

There is a continuum of symmetric monopolistically competitive producers for the consumption goods, each producing a differentiated variety $z \in [0,1]$.[33] The production function of firm z is $y_t(z) = X_t(z)$, where $X_t(z)$ is the amount of aggregate intermediate goods employed in the consumption goods production process of firm z.

The consumption basket C_t takes the constant elasticity of substitution (CES) form: $C_t = \left[\int_0^1 c_t(z)^{\frac{\gamma-1}{\gamma}} dz\right]^{\frac{\gamma}{\gamma-1}}$, where $1 < \gamma \ll \infty$ is the elasticity of substitution across the consumption goods, $c_t(z)$ is the demand for individual firm z's goods. It follows from the CES consumption basket assumption that the household demand for firm z's goods is $c_t(z) = [\frac{p_t(z)}{P_t^C}]^{-\gamma} C_t$, where $p_t(z)$ is the price of firm z's good, and $P_t^C \equiv [\int_0^1 p_t(z)^{1-\gamma} dz]^{\frac{1}{1-\gamma}}$ is the ideal consumer price index (CPI).

We assume there is price inertia in the consumption goods sector. More specifically, we follow Rotemberg (1982) and Bilbiie et al. (2008) to assume that the consumption goods producer z has to pay a price setting cost of the form $pac_t(z) = \frac{\eta}{2}\left[\Phi_t(z) - 1\right]^2 p_t(z) y_t(z)$, where $\eta \in [0,\infty)$ and $\Phi_t(z) \equiv \frac{p_t(z)}{p_{t-1}(z)}$ is firm z's gross price inflation. Following Erceg et al. (2000), we assume that the government provides a subsidy $\tau_c = \frac{1}{\gamma-1}$ to the firm's output so that the distortion from monopolistic competition in the consumption goods sector is eliminated. Therefore, firm z's periodic real profit is

$$\begin{aligned} m_t^c(z) &= \{(1+\tau_c)p_t(z)y_t(z) - P_t^M X_t(z) - \tfrac{\eta}{2}\left[\Phi_t(z) - 1\right]^2 p_t(z) y_t(z)\}/P_t^C \\ &= \{(1+\tau_c)p_t(z) - P_t^M - \tfrac{\eta}{2}\left[\Phi_t(z) - 1\right]^2 p_t(z)\} y_t(z)/P_t^C, \end{aligned}$$

[33]The fixed number of varieties has been normalized to unity.

where P_t^M is the price index of the aggregate intermediate goods, and the second equality comes from our specification of the consumption goods production function.

One purpose of introducing the subsidy in the model is to facilitate comparison of the quantitative results with those of Bilbiie et al. (2008) since they introduce a government subsidy to eliminate the distortion from monopolistic competition in their model. Moreover, if the distortion from monopolistic competition is eliminated, the central bank's monetary policy only has to concern about frictions from nominal rigidities and the financial sector. Note that given nominal rigidities in our model, the distortion from monopolistic competition can be completely eliminated by the subsidy only in the steady state. The interactions between monopolistic competition and nominal rigidities still affect the transitional dynamics.

Firm z chooses a price level to maximize the net present value (NPV) of the profit flows $E_t\Sigma_{s=t}^{\infty}[\Lambda_{t,s}m_s^c(z)]$, where $\Lambda_{t,s}\equiv\beta^{s-t}(U_{Cs}/U_{Ct})$ is the stochastic discount factor, β is the subjective discount factor, and U_{Cs} is the marginal utility of consumption in period s. Following Bilbiie et al. (2008), we interpret the real price setting cost as the amount of marketing materials needed to set the price and assume that the basket of the marketing materials has the same composition as the consumption basket. Therefore, the demand function for firm z's goods is $y_t(z)=[\frac{p_t(z)}{P_t^C}]^{-\gamma}Y_t$. Maximizing the NPV of firm z's profit flows subject to the demand function, we obtain the optimal pricing condition for the consumption good producer firm z: $p_t(z)=\mu_t(z)P_t^M$, where $\mu_t(z)$ is the markup over marginal cost defined as

$$\mu_t(z)=\frac{\gamma}{(\gamma-1)\left\{1+\tau_c-\frac{\eta}{2}[\Phi_t(z)-1]^2\right\}+\Gamma_t(z)},$$

$$\Gamma_t(z)\equiv\eta\left\{\Phi_t(z)[\Phi_t(z)-1]-\beta E_t\left[\frac{U_{Ct+1}}{U_{Ct}}\frac{Y_{t+1}}{Y_t}\left(\Phi_{t+1}^C\right)^{\gamma-1}\Phi_{t+1}(z)^{2-\gamma}\left(\Phi_{t+1}(z)-1\right)\right]\right\},$$

where $\Phi_t^C\equiv\frac{P_t^C}{P_{t-1}^C}$ is the gross consumer price inflation rate. Note that in the steady state with no price adjustment, the markup is one. This result obtains because the monopolistic distortion is eliminated by the production subsidy.

Imposing symmetry, it is easy to see that the producer price inflation rate of the consumption goods sector is also the CPI inflation rate. More specifically, when producers are symmetric, the aggregate pricing equation of the consumption goods sector reduces to $P_t^C=[\int_0^1 p_t^{1-\gamma}dz]^{\frac{1}{1-\gamma}}=p_t$, where $p_t=p_t(z)$ is the average producer price in the consumption goods sector.

Intermediate goods producers

Similar to the consumption goods production sector, the intermediate goods production sector also features monopolistic competition. However, different from the consumption goods sector, we assume that the number of varieties of the intermediate goods can change over time due to free entry and exit. More specifically, we assume that there is a continuum of intermediate goods producers, each producing a different variety $\omega \in \Omega$. A basket of the intermediate goods is produced according to $X_t = \left[\int_\Omega x_t(\omega)^{\frac{\epsilon-1}{\epsilon}} d\omega\right]^{\frac{\epsilon}{\epsilon-1}}$, where $1 < \epsilon \ll \infty$ is the elasticity of substitution across intermediate goods. Hence, the individual intermediate goods firm ω's demand function is $x_t(\omega) = [\frac{p_t^m(\omega)}{P_t^M}]^{-\epsilon} X_t$, where $p_t^m(\omega)$ is the price of firm ω's good, $P_t^M = [\int_\Omega p_t^m(\omega)^{1-\epsilon}]^{\frac{1}{1-\epsilon}}$ is the aggregate price index of the intermediate goods.

Following Ghironi and Melitz (2005), we assume that the production function of the intermediate goods firm ω is $x_t(\omega) = \omega Z_t l_t(\omega)$, where $l_t(\omega)$ is the labor input for production, Z_t is the stochastic aggregate productivity level, ω is the individual productivity level which is drawn after entry and remains fixed thereafter. Hence, the unit labor cost for intermediate goods production is $w_t/\omega Z_t$, where w_t is the aggregate real wage rate. We assume that there is no price adjustment cost in the intermediate goods sector. Similar to the consumption goods sector, there is an output subsidy $\tau_m = \frac{1}{\epsilon-1}$ so that distortion from monopolistic competition is eliminated. Given those assumptions, the real gross profit function of the intermediate goods firm ω is

$$
\begin{aligned}
m_t(\omega) &= [(1+\tau_m)p^m(\omega)/P_t^C - w_t/\omega Z_t]x_t(\omega) \\
&= [(1+\tau_m)p_t^m(\omega) - p_t^m(\omega)]x_t(\omega)/P_t^C \\
&= \tau_m \tfrac{P_t^M}{P_t^C}[\tfrac{p_t^m(\omega)}{P_t^M}]^{1-\epsilon} X_t,
\end{aligned}
$$

where the second equation follows from firms setting their price equal to the marginal cost, $p_t^m(\omega) = \frac{w_t P_t^C}{\omega Z_t}$, due to the subsidy; the third equation is the result of substituting in the demand function $x_t(\omega) = [\frac{p_t^m(\omega)}{P_t^M}]^{-\epsilon} X_t$. Obviously, firms with a higher individual productivity ω earn more profit.

Aggregation, and entry and exit of intermediate goods producers

To enter the market, the intermediate goods producers have to pay a sunk cost. The sunk cost is composed of two parts. One part is an amount of effective labor cost ($\frac{w_t}{Z_t} f^{ew}$) which is covered by the firm's own money.[34] The other part is the cost of purchasing a fixed amount (f^e) of aggregate consumption goods covered by loans from financial intermediaries.[35] The loan is then repaid by installments in each period. As we shall see in subsection 3.2.2, the periodic installment (f_t) is predetermined and unaffected by an individual firm's productivity. This specification suggests that the probability that firm ω is able to pay the full amount of installment is higher when its individual productivity ω is higher. Therefore, there is a cutoff individual productivity level ω_t^*, which satisfies $m_t(\omega_t^*) = f_t$. Note that we add a time subscript to the cutoff individual productivity as it varies with aggregate productivity.

We assume that firms that fail to pay the full amount of the installment will go bankrupt. This assumption implies that the bankruptcy law imposes a strict solvency constraint on the borrowers so that all defaulting borrowers will be forced to go bankrupt even if some of them may be able to repay the debt in the future, once the aggregate economic situation has become more favorable. In practice, bankruptcy laws differ across countries. For example, bankruptcy laws in the UK are much stricter than in the US. In the US, there is Chapter 11, which allows the firms in financial distress to reorganize and continue to operate afterward. We do not model this for tractability reasons. However, the existence of a soft budget constraint is likely to deter the entry of entrepreneurs and worsen the borrower pool faced by financial intermediaries, since keeping more firms in the market could reduce the expected profit flows for an entering entrepreneur. Moreover, a soft budget constraint could encourage gambling since it gives gamblers a better chance to survive longer and benefit more from taking the gamble. In this sense, introducing a soft budget constraint may strengthen rather than weaken the results of the current model.

We further assume that there is limited liability which means that the firms do not have to pay an amount more than their profits to the lender when they go bankrupt. Therefore, those firms which expect to earn profits less than the installment will exit the

[34] More precisely, it is indirectly covered by the households owning the firms.

[35] This technical assumption simplifies aggregation because the principal amount of the debt is the same for firms entering the market in different periods under this assumption.

market without production since they can earn nothing from producing. This is different from the traditional financial accelerator model (Bernanke et al., 1999) in which the firms' current period profit is modeled as a collateral for the loan. However, it is consistent with the theoretical model of Dell'Ariccia and Marquez (2006), which suggests that if the number of new loan applications is sufficiently large, financial intermediaries will optimally choose not to screen the borrowers and require no collateral from them.

With these assumptions, we can aggregate the intermediate goods production sector in the same way as Melitz (2003) and Ghironi and Melitz (2005) have done. More specifically, we assume that the intermediate goods producers draw their individual productivity levels from a Pareto distribution $G(\omega) = 1 - (1/\omega)^k$ over $[1, \infty)$. Then, an average productivity level defined as

$$\begin{aligned}\omega_t^m &\equiv [\tfrac{1}{1-G(\omega_t^*)} \int_{\omega_t^*}^{\infty} \omega^{\epsilon-1} dG(\omega)]^{1/(\epsilon-1)} \\ &= [k/(k-\epsilon+1)]^{1/(\epsilon-1)} \omega_t^*\end{aligned}$$

can summarize all the information on the individual productivity distributions relevant for all macroeconomic variables. Essentially, the intermediate goods producer block of our model with N_t firms with heterogeneous productivity is isomorphic to one where N_t representative firms with productivity ω_t^m produce the intermediate goods. Particularly, we have $P_t^M = N_t^{\overline{\frac{1}{1-\epsilon}}} p_t^m(\omega_t^m)$, which is a result of Melitz (2003).

Following Ghironi and Melitz (2005) and Bilbiie et al. (2008), we assume that there is a time-to-build lag such that the firms start producing only one period after paying the sunk cost. Firms with an individual productivity level higher than ω_t^* will not go bankrupt, so the firm survival rate in period t is $\theta_t \equiv 1 - G(\omega_t^*) = (1/\omega_t^*)^k$. Therefore, an entering firm's average value in period t is $v_t = E_t \Sigma_{s=t+1}^{\infty} \{\Lambda_{t,s}(\prod_{t+1}^{s} \theta_s)[m_s(\omega_s^m) - f_s]\}$. Free entry in the intermediate goods production sector requires that the average value of the firm equals the sunk cost paid with own funds:

$$v_t = \tfrac{w_t}{Z_t} f^{ew}.$$

Denoting the number of new entrants in the intermediate goods sector by N_t^e, we get the dynamic equation for the number of producing firms: $N_t = \theta_t(N_{t-1} + N_{t-1}^e)$.

3.2.2. Financial intermediation

In each period, there are a number (N_t^r) of investors who come to the financial intermediaries for funding. A proportion $\phi_t = N_t^e/N_t^r$ of those investors are entrepreneurs who will invest the borrowed money in the intermediate goods sector to start their business. The other $(1-\phi_t)N_t^r$ investors are gamblers who will invest the borrowed money on a pure gambling asset of which the supply is fixed for each period. The loan from the intermediaries is paid back by a periodic installment (f_s) from one period after the borrowing. We introduce the one-period lag here to allow for a time-to-build lag in the real sector. We assume that the financial intermediaries do not screen out gamblers from the borrower pool. As a result, the borrowing amount and periodic repayment will be the same for both entrepreneurs and gamblers. Therefore, the borrowed money of a gambler in period t is f^e which is equal to the part of sunk cost of an intermediate goods producer covered by a loan from the financial intermediaries.

One period after purchase, the buyer of the gambling asset can participate in a lottery, which gives a payoff g with probability λ and a payoff zero with probability $1-\lambda$. Conditional on winning the lottery, the owner of the gambling asset can participate in the same lottery again in the next period. The gambler can keep participating in the lottery until he fails to win the lottery. Denote the real gambling asset price by p_t^r. Then, the number of gambling asset bought by a gambler is $\frac{f^e}{p_t^r}$. Similar to the entrepreneurs, gamblers will go bankrupt if they cannot pay the full amount of installment, and their profit is zero when bankrupt due to protection by the limited liability law. Therefore, the expected payoff of a gambler is $E_t\{\Sigma_{s=t+1}^{\infty}[\Lambda_{t,s}\lambda^{s-t}[\prod_{t+1}^{s} Prob(g\frac{f^e}{p_t^r} \geq f_s)](g\frac{f^e}{p_t^r} - f_s)]\}$, where $Prob(x)$ denotes the probability of event x.[36] Assuming that the gamblers have to pay an entrance fee[37] (f^g) for the gambling market with own money, we can write the free entry condition of the gambling market as $E_t\{\Sigma_{s=t+1}^{\infty}[\Lambda_{t,s}\lambda^{s-t}[\prod_{t+1}^{s} Prob(g\frac{f^e}{p_t^r} \geq f_s)](g\frac{f^e}{p_t^r} - f_s)]\} = f^g$.

[36] Here the analysis is simplified by assuming that the investors cannot sell the assets in a secondary market. In other words, they are locked up after purchasing. Ofek and Richardson (2003) provide evidence that lockup agreements are responsible for the buildup of the internet bubble. In practice, the gamblers could be the existing business owners who starts excessively risky new projects with easy money from the financial intermediaries. Typically, selling of the projects involves very high liquidation costs. Therefore, it is reasonable to assume no resale of those assets. The lockup assumption is a very stringent form of short sale constraint. Our intuition is that a less stringent form of short sale constraint should be enough to keep the bubble. Kocherlakota (2008) shows that short sale constraints can arise endogenously. Hence, the model's results could be more general.

[37] This could be the searching cost for the gambling opportunity, for instance.

Assuming that g is large enough so that $g\frac{f^e}{p_t^r} \geq f_s$ always holds, the above equation reduces to $E_t\{\Sigma_{s=t+1}^{\infty}[\Lambda_{t,s}\lambda^{s-t}(g\frac{f^e}{p_t^r} - f_s)]\} = f^g$ which implies that the real asset price is

$$p_t^r = \frac{f^e g E_t[\Sigma_{s=t+1}^{\infty}(\Lambda_{t,s}\lambda^{s-t})]}{f^g + E_t[\Sigma_{s=t+1}^{\infty}(\Lambda_{t,s}\lambda^{s-t}f_s)]}.$$

Following Allen and Gale (2000), we define the fundamental value of the gambling asset as the NPV of the returns from the gambling asset when the gamblers have to buy it with their own money. More specifically, the fundamental value is $gE_t[\Sigma_{s=t+1}^{\infty}(\Lambda_{t,s}\lambda^{s-t})]$. It is easy to see that when $\frac{f^e}{f^g + E_t[\Sigma_{s=t+1}^{\infty}(\Lambda_{t,s}\lambda^{s-t}f_s)]} > 1$, the real asset price is larger than its fundamental value. This reflects the idea of Allen and Gale (2000) that excessive risk-taking behavior induced by the limited liability law can create asset price bubbles. More specifically, for bubbles to exist, entry into the real sector must be more difficult than entry into the gambling sector, i.e., $f^e > f^g$ must hold. Additionally, the NPV of expected repayments from the gambler ($E_t[\Sigma_{s=t+1}^{\infty}(\Lambda_{t,s}\lambda^{s-t}f_s)]$) must be relatively small compared to the amount borrowed (f^e). This suggests that lending to a gambler cannot be good business for the financial intermediaries if there is a bubble in the gambling asset price. However, even in this case, the financial intermediaries may still be willing to lend to loan applicants because expected returns from lending to the entrepreneurs could cover the expected losses from lending to the gamblers.

To facilitate impulse response analysis later, we define $DV_{1t} \equiv E_t[\Sigma_{s=t+1}^{\infty}(\Lambda_{t,s}\lambda^{s-t})]$ and $DV_{2t} \equiv E_t[\Sigma_{s=t+1}^{\infty}(\Lambda_{t,s}\lambda^{s-t}f_s)]$. These two definitions can be written in recursive forms: $DV_{1t} = E_t(\Lambda_{t,t+1}\lambda) + E_t(\Lambda_{t,t+1}\lambda DV_{1t+1})$, $DV_{2t} = E_t(\Lambda_{t,t+1}\lambda f_{t+1}) + E_t(\Lambda_{t,t+1}\lambda DV_{2t+1})$. Denote the fundamental value of the gambling asset by fv_t and define the bubble size as

$$bb_t \equiv \frac{f^e}{f^g + E_t[\Sigma_{s=t+1}^{\infty}(\Lambda_{t,s}\lambda^{s-t}f_s)]},$$

then $fv_t = gDV_{1t}$, $bb_t = \frac{f^e}{f^g + DV_{2t}}$, and $p_t^r = bb_t fv_t$.

Rather than explicitly modeling the pricing behavior of the financial intermediaries, we adopt the reduced-form specification of Chowdhury et al. (2006):

$$\hat{r}_{t+1}^b = (1 + \phi_r)\hat{r}_t^m,$$

where a hat over a variable denotes log-deviation from its steady state, r_t^m is the real gross money market interest rate, and r^b_{t+1} is the real gross loan rate which satisfies $(r^b_{t+1} - 1)f^e = f_{t+1}$. Note that we use the beginning of the period timing for r^b_{t+1}, so the above equation actually describes the evolution of current period credit spread. $(1 + \phi_r)$ captures the interest rate pass-through.[38] When $\phi_r = -1$, the interest rate passthrough is zero.

3.2.3. Labor market structure and wage setting

Following Erceg et al. (2000), we assume that there is monopolistic competition in the labor market. Each household $j \in (0, 1)$ supplies a differentiated labor type $H(j)$ to the market and the aggregate labor demand is $L_t = \left[\int_0^1 H_t(j)^{\frac{\epsilon_w - 1}{\epsilon_w}} dj\right]^{\frac{\epsilon_w}{\epsilon_w - 1}}$, where $1 < \epsilon_w \ll \infty$. The demand for each labor type j is then

$$H_t(j) = \left[\frac{W_t(j)}{W_t}\right]^{-\epsilon_w} L_t,$$

where $W_t(j)$ is the nominal wage of the jth household, $W_t \equiv \left[\int_0^1 W_t(j)^{1-\epsilon_w} dj\right]^{\frac{1}{1-\epsilon_w}}$ is the aggregate nominal wage rate.

There is nominal wage rigidity. More specifically, a household can reset its nominal wage rate with a fixed probability $1 - \eta_w$ in each period, where $\eta_w \in (0, 1)$. The nominal wage rate of those who cannot reoptimize face the wage rate from the last period, that is, $W_t(j) = W_{t-1}(j)$ if j cannot reset its wage. The real wage is defined as $w_t(j) \equiv \frac{W_t(j)}{P^C_t}$.

3.2.4. Households

In each period, the household j gets a working salary (in real terms) from the firms $w_t(j)H_t(j)$. Following Erceg et al. (2000), we assume that the government subsidizes the workers with a subsidy rate $\tau_l = \frac{1}{\epsilon_w - 1}$ to eliminate monopolistic distortion from the labor market, so the actual real labor income is $(1 + \tau_l)w_t(j)H_t(j)$. The households also get profits from the firms. More specifically, they get profits m^c_t from the consumption goods producers and profits $N_t[m_t(\omega^m_t) - f_t]$ from the intermediate goods producers. Here, we make use of the result of Melitz (2003) that the firm with the average productivity ω^m_t earns the average profit in the market. Note that the total profit in the intermediate

[38] See Ravenna and Walsh (2006), Nabar et al. (1993), Sander and Kleimeier (2004) for summaries of theoretical discussions.

goods sector net of subsidy is zero because the price is set to marginal cost when there is a production subsidy. The total profit in the consumption goods sector net of subsidy is $(m_t^c - \tau_c Y_t)$. In the steady state without price adjustment it is also zero because the price markup is driven to one by the subsidy. However, the markup can deviate from one if there is nominal price adjustment. In that case, $(m_t^c - \tau_c Y_t)$ will be different from zero. Besides the labor income and profit dividends from the firms, the households also get the repayment of their deposits from the financial intermediaries $r_t^m S_t$, where S_t is the amount of deposits in period t.

Because of nominal wage rigidity, it is uncertain whether the household j could reoptimize its wage. This could generate discrepancy in labor incomes between those who can reset their wage rates and those who cannot. Hence, the decision on saving and spending could differ across households. Following Christiano et al. (2005), we assume that there are short-term securities with payoffs contingent on whether households can reset their nominal wage. This ensures that the households are homogeneous in terms of consumption, investment and deposit, though they are heterogeneous in terms of wage setting and labor supply. Therefore, we can treat the household j as a representative household in terms of consumption, investment, deposit and claims on profit.

In sum, the household j's wealth in each period is $(1+\tau_l)w_t(j)H_t(j)+m_t^c+N_t[m_t(\omega_t^m)-f_t] + r_t^m S_t + A_t(j)$, where $A_t(j)$ is the payoff from the state-contingent securities. The households use their wealth to consume C_t, invest $N_t^e v_t$ to build new production lines in the intermediate goods sector, deposit S_{t+1} to the financial intermediaries and pay a lump-sum tax T_t^L (defined in real terms) to the government. Therefore, the household budget constraint is $C_t + N_t^e v_t + S_{t+1} + T_t^L = (1+\tau_l)w_t(j)H_t(j) + m_t^c + N_t[m_t(\omega_t^m) - f_t] + r_t^m S_t + A_t(j)$.

The household chooses deposits (S_{t+1}) and labor supply (H_t) to maximize its expected intertemporal utility $E_t \Sigma_{s=t}^{\infty} \beta^{s-t} U(C_s, H_s)$, where β is the subjective discount factor, $U(C_s, H_s)$ is the periodic utility function in period s, C_s is aggregate consumption, H_s is labor supply. Christiano et al. (2005) suggest that it is necessary to model habit formation to capture the hump-shaped response of consumption to the monetary policy shock. Following them, we model habit formation as the dependence of the current period's utility on the last period's consumption. More specifically, we have $U(C_s, H_s) = ln(C_s - bC_{s-1}) - \frac{\chi H_s^{1+1/\phi_l}}{1+1/\phi_l}$, where b is the parameter governing the relative importance of habit formation, χ is a positive parameter, and ϕ_l is the Frisch labor supply elasticity. The maximization problem gives the first-order condition (FOC) for deposit:

$$U_{Ct} = \beta E_t(r^m_{t+1} U_{Ct+1}),$$

where $U_{Ct} \equiv (C_t - bC_{t-1})^{-1} - \beta b[E_t(C_{t+1}) - bC_t]^{-1}$ is the marginal utility of consumption. The FOC for deposit suggests that the marginal disutility of giving up current consumption must be equal to the expected utility gain from the corresponding increase in next period's consumption.

The households that can reset their wage rate choose the reset wage rate (W^*_t) to maximize $E_t \Sigma^{\infty}_{s=t} (\beta\eta_w)^{s-t} U(C_{s|t}, H_{s|t}(j))$, where $X_{s|t}$ denotes the value of variable X in period s for the households which last reset their wage rates in period t. The corresponding FOC is:

$$\Sigma^{\infty}_{s=t} (\beta\eta_w)^{s-t} E_t \left\{ H_{s|t}(j) \left[U_{Cs|t} \frac{W^*_t}{P^C_s} - \chi H_{s|t}(j)^{1/\phi_l} \right] \right\} = 0.$$

It determines the optimal reset wage rate and labor supply. In case there is no nominal wage rigidity, the FOC for labor supply is $U_{Ct} \frac{W_t(j)}{P^C_s} = \chi H_t(j)^{1/\phi_l}$, which suggests that the marginal disutility from working must be equal to the utility gain from the corresponding increase in consumption. Note that this FOC is the same as the one in a perfectly competitive labor market, as the distortion from labor market monopoly power is eliminated by the labor subsidy.

3.2.5. Market clearing and aggregate accounting

Consumption goods market clearing requires that the output of consumption goods equals its demand from consumption (C_t), investment ($N^e_t f^e$), and marketing (PAC_t), where $PAC_t \equiv pac_t(z)/P^C_t$ is both the average and aggregate real price setting cost since the number of consumption goods producers is normalized to one and the consumption goods producers are symmetric. More specifically, we have $Y_t = C_t + N^e_t f^e + PAC_t$. Substituting the definition of $pac_t(z)$ into the market clearing condition, we get $Y_t = C_t + N^e_t f^e + \frac{\eta}{2} (\Phi_t - 1)^2 p_t y_t / P^C_t$, where we have omitted the index z by applying the symmetry assumption across consumption goods producers. Combining this equation with the demand function of the consumption good producer $y_t = (\frac{p_t}{P_t^C})^{-\gamma} Y_t$, we obtain

$$C_t + N_t^e f^e = [1 - \frac{\eta}{2} (\Phi_t - 1)^2] p_t^{1-\gamma} (P_t^C)^{\gamma-1} Y_t.$$

Each consumption goods producer demands $X_t(z) = y_t(z) = [\frac{p_t(z)}{P_t^C}]^{-\gamma} Y_t$ amount of aggregate intermediate goods and the number of consumption goods producers is normalized to one, so $[\frac{p_t(z)}{P_t^C}]^{-\gamma} Y_t$ is the total demand for the aggregate intermediate goods. Intermediate goods market clearing then requires $[\frac{p_t(z)}{P_t^C}]^{-\gamma} Y_t = X_t$.

The government budget constraint requires that the tax revenue equals the sum of all subsidies, that is, $T_t^L = \tau_l w_t L_t + \tau_c Y_t + \tau_m P_t^M X_t / P_t^C$. Here, we use the result that total production subsidies (in real terms) to the consumption goods sector and intermediate goods sector are $\int_0^1 \tau_c p_t(z) y_t(z) dz / P_t^C = \tau_c Y_t$ and $\int_\Omega \tau_m p_t^m(\omega) x_t(\omega) d\omega / P_t^C = \tau_m P_t^M X_t / P_t^C$, respectively. Combining the government and household budget constraint, we get the aggregate accounting identity $C_t + N_t^e v_t + S_{t+1} = w_t L_t + m_t^c - \tau_c Y_t - N_t f_t + r_t^m S_t$.[39]

Gambling asset demand is equal to the number of gamblers multiplied by the per gambler purchase, that is, $(N_t^r - N_t^e) \frac{f^e}{P_t^r}$. Denote the periodically fixed supply of the gambling asset by GS. Then the gambling market clearing condition is $(N_t^r - N_t^e) \frac{f^e}{P_t^r} = GS$. Finally, loan market equilibrium requires total saving equal to total lending: $S_{t+1} = N_t^r f^e$.

3.2.6. Monetary policy rules

Following Bilbiie et al. (2008), we define the gross real money market interest rate by $r_t^m \equiv i_{t-1}^m / \Phi_t^C$, where i_t^m is the gross nominal money market interest rate. The nominal money market interest rate is set by the central bank according to a specific feedback rule. We consider three different monetary policy rules in our analysis. The first two rules involve interest rate smoothing. One of those two rules does not require the interest rate to react to output while the other one does. More specifically, one rule has the following

[39] The aggregate payoff from the state-contingent securities is zero.

form[40]

$$\hat{i}_t^m = \rho \hat{i}_{t-1}^m + (1-\rho)(1.5\pi_{t+1}),$$

while the other rule is

$$\hat{i}_t^m = \rho \hat{i}_{t-1}^m + (1-\rho)(1.5\pi_{t+1} + 0.1\hat{y}_{t+1}^a),$$

where the real GDP level y_t^a is defined as the sum of consumption (C_t) and investment ($N_t^e v_t$), and the smoothing parameter ρ is set to 0.8 so that the first interest rate smoothing rule is identical to the one used by Bilbiie et al. (2008), while the second rule is identical to the one used by Christiano et al. (2005). The third monetary policy rule is a forward-looking Taylor rule without interest rate smoothing:

$$\hat{i}_t^m = 1.5\pi_{t+1} + (0.5/4)\hat{y}_{t+1}^a,$$

where the 0.5 coefficient of Taylor's original specification (Taylor, 1993) is divided by 4 since the annualized inflation and interest rate in Taylor's original specification are replaced by quarterly inflation and interest rate in the current chapter.

Note that $\hat{y}_t^a$ is the deviation of real GDP from its flexible-price steady-state level. It is equal to the theoretical output gap in case there is no technology shock, but will diverge from the theoretical output gap if there is a technology shock to the economy. However, as noted by Woodford (2003), the widely used empirical output gap estimated as the deviation of output from a smooth trend can be very different from the theoretical output gap. Neiss and Nelson (2005) estimate the theoretical output gap for the US, UK, and Australia and find that the empirical output gap estimates from detrending methods are very different from the theoretical output gap. Further, troughs in the HP-filtered output gap accord well with the recessions documented by the National Bureau of Economic Research (NBER) (Rudd and Whelan, 2007), which suggests that by targeting the output gap generated by detrending methods such as the HP filter, central banks are actually targeting output fluctuations ($\hat{y}_t^a$) rather than the theory-consistent output gap.

[40]Note that $\hat{i}_t^m$ is the log-deviation of the gross nominal money market interest rate from its steady state. Therefore, $\hat{i}_t^m = log(1+i_t^n) - log(1+\bar{i}^n) \doteq i_t^n - \bar{i}^n$, where i_t^n is the net nominal money market interest rate and $\bar{i}^n$ is its steady state level. We can replace $\hat{i}_t^m$ by $i_t^n - \bar{i}^n$ in the policy rules. However, working with the gross interest rate simplifies log-linearization of the model.

3.2.7. Model summary

Table 3.1 summarizes the main equations of the model.[41] The infinite sum V_t defined in the text is rewritten in recursive form in the table. The real profit equation of the consumption goods sector in the table is the result of substituting the pricing equation and demand function of the consumption goods sector into the real profit function in the text. The model can be simplified by using the aggregate pricing equation of the consumption goods sector $P_t^C = p_t$ to substitute for P_t^C in the other equations of the system. Moreover, the price levels $p_t, P_t^M, p_t^m(\omega_t^m), p_t^m(\omega_t^*)$ are not stationary in the model. To simulate the model, we have to transform it to make all the variables in the model stationary. This is done by defining the real price of aggregate intermediate goods by $q_t = P_t^M/p_t$ and using it to substitute for the nominal price levels in the model. The transformed model is summarized in Table 3.2. Note that we follow Bilbiie et al. (2008) by using beginning of the period timing, so $r_{t+1}^b, f_{t+1}, S_{t+1}$ are actually determined in period t. The model can be closed by specifying a process of the exogenous variable Z_t and the parameters: β, b, f^e, f^{ew}, γ, ϵ, ϵ_w, η, η_w, k, χ, ϕ_l, ϕ_r, GS, λ, f_g, which we will do in the next section.

[41] We do not include the specification of the monetary policy rule in the summary table to save space.

Table 3.1: Model Summary

Consumption goods sector:	
Pricing	$p_t = \mu_t P_t^M$
Markup	$\mu_t = \frac{\gamma}{(\gamma-1)\left[1+\tau_c-\frac{\eta}{2}(\Phi_t-1)^2\right]+\Gamma_t(z)}$
Real profit	$m_t^c = \{1+\tau_c-1/\mu_t-\frac{\eta}{2}\left[\Phi_t-1\right]^2\}p_t^{1-\gamma}(P_t^C)^{\gamma-1}Y_t$
Aggregate pricing	$P_t^C = p_t$
CPI inflation	$\Phi_t^C = P_t^C/P_{t-1}^C$
Producer price inflation	$\Phi_t = p_t/p_{t-1}$
Intermediate goods sector:	
Average individual productivity	$\omega_t^m = [k/(k-\epsilon+1)]^{1/(\epsilon-1)}\omega_t^*$
Pricing	$p_t^m(\omega_t^m) = \frac{w_t P_t^C}{\omega_t^m Z_t}$
	$p_t^m(\omega_t^*) = \frac{w_t P_t^C}{\omega_t^* Z_t}$
Real Profit	$m_t(\omega_t^m) = \tau_m \frac{P_t^M}{P_t^C}[\frac{p_t^m(\omega_t^m)}{P_t^M}]^{1-\epsilon}X_t$
	$m_t(\omega_t^*) = \tau_m \frac{P_t^M}{P_t^C}[\frac{p_t^m(\omega_t^*)}{P_t^M}]^{1-\epsilon}X_t$
Aggregate pricing	$P_t^M = N_t^{\frac{1}{1-\epsilon}} p_t^m(\omega_t^m)$
Firm value	$v_t = E_t\{\beta\frac{U_{Ct+1}}{U_{Ct}}\theta_{t+1}[m_{t+1}(\omega_{t+1}^m)-f_{t+1}]\} + E_t\left(\beta\frac{U_{Ct+1}}{U_{Ct}}\theta_{t+1}v_{t+1}\right)$
Free entry	$v_t = \frac{w_t}{Z_t}f^{ew}$
Cutoff condition	$m_t(\omega_t^*) = f_t$
Survival rate	$\theta_t = (1/\omega_t^*)^k$
Number of firms	$N_t = \theta_t(N_{t-1}+N_{t-1}^e)$
Financial sector:	
Fundamental value(gamble)	$fv_t = gDV_{1t}$
Bubble size	$bb_t = \frac{f^e}{f^g+DV_{2t}}$
Asset price(gamble)	$p_t^r = bb_t fv_t$
Gambling asset market clearing	$(N_t^r - N_t^e)\frac{f^e}{P_t^r} = GS$
Proportion of entrepreneur	$\phi_t = N_t^e/N_t^r$
Definition of loan rate	$(r_{t+1}^b - 1)f^e = f_{t+1}$
Evolution of loan rate	$\hat{r}_{t+1}^b = (1+\phi_r)\hat{r}_t^m$
Households:	
Euler equation (deposit)	$U_{Ct} = \beta E_t(r_{t+1}^m U_{Ct+1})$
Labor supply	$\Sigma_{s=t}^{\infty}(\beta\eta_w)^{s-t}E_t\left\{H_{s\|t}(j)\left[U_{Cs\|t}\frac{W_t^*}{P_s^C} - \chi H_{s\|t}(j)^{1/\phi_l}\right]\right\} = 0$
Market clearing:	
Consumption goods market	$[1-\frac{\eta}{2}(\Phi_t-1)^2]p_t^{1-\gamma}(P_t^C)^{\gamma-1}Y_t = C_t + N_t^e f^e$
Intermediate goods market	$[\frac{p_t}{P_t^C}]^{-\gamma}Y_t = X_t$
Loan market	$S_{t+1} = N_t^r f^e$
Aggregate accounting	$C_t + N_t^e v_t + S_{t+1} = w_t L_t + m_t^c - \tau_c Y_t - N_t f_t + r_t^m S_t$

Notes: $DV_{1t} \equiv E_t[\Sigma_{s=t+1}^{\infty}(\Lambda_{t,s}\lambda^{s-t})] = E_t(\Lambda_{t,t+1}\lambda) + E_t(\Lambda_{t,t+1}\lambda DV_{1t+1})$.
$DV_{2t} \equiv E_t[\Sigma_{s=t+1}^{\infty}(\Lambda_{t,s}\lambda^{s-t}f_s)] = E_t(\Lambda_{t,t+1}\lambda f_{t+1}) + E_t(\Lambda_{t,t+1}\lambda DV_{2t+1})$.
$\Gamma_t(z) \equiv \eta\left\{\Phi_t(z)[\Phi_t(z)-1] - \beta E_t\left[\frac{U_{Ct+1}}{U_{Ct}}\frac{Y_{t+1}}{Y_t}\left(\Phi_{t+1}^C\right)^{\gamma-1}\Phi_{t+1}(z)^{2-\gamma}\left(\Phi_{t+1}(z)-1\right)\right]\right\}$.
$U_{Ct} \equiv (C_t - bC_{t-1})^{-1} - \beta b[E_t(C_{t+1}) - bC_t]^{-1}$ is the marginal utility of consumption.

Table 3.2: Transformed Model Summary

Consumption goods sector:	
Pricing	$\mu_t q_t = 1$
Markup	$\mu_t = \frac{\gamma}{(\gamma-1)\left[1+\tau_c-\frac{\eta}{2}(\Phi_t-1)^2\right]+\Gamma_t(z)}$
Real profit	$m_t^c = \{1+\tau_c - 1/\mu_t - \frac{\eta}{2}\left[\Phi_t - 1\right]^2\}Y_t$
Intermediate goods sector:	
Average individual productivity	$\omega_t^m = [k/(k-\epsilon+1)]^{1/(\epsilon-1)}\omega_t^*$
Real Profit	$m_t(\omega_t^m) = \tau_m(\frac{w_t}{\omega_t^m Z_t})^{1-\epsilon} q_t^\epsilon X_t$
	$m_t(\omega_t^*) = \tau_m(\frac{w_t}{\omega_t^* Z_t})^{1-\epsilon} q_t^\epsilon X_t$
Aggregate pricing	$q_t = \frac{w_t}{\omega_t^m Z_t} N_t^{\frac{1}{1-\epsilon}}$
Firm value	$v_t = E_t\{\beta \frac{U_{Ct+1}}{U_{Ct}}\theta_{t+1}[m_{t+1}(\omega_{t+1}^m) - f_{t+1}]\} + E_t\left(\beta\frac{U_{Ct+1}}{U_{Ct}}\theta_{t+1}v_{t+1}\right)$
Free entry	$v_t = \frac{w_t}{Z_t} f^{ew}$
Cutoff condition	$m_t(\omega_t^*) = f_t$
Survival rate	$\theta_t = (1/\omega_t^*)^k$
Number of firms	$N_t = \theta_t(N_{t-1} + N_{t-1}^e)$
Financial sector:	
Fundamental value(gamble)	$fv_t = gDV_{1t}$
Bubble size	$bb_t = \frac{f^e}{f^g + DV_{2t}}$
Asset price(gamble)	$p_t^r = bb_t fv_t$
Gambling asset market clearing	$(N_t^r - N_t^e)\frac{f^e}{P_t^r} = GS$
Proportion of entrepreneur	$\phi_t = N_t^e / N^r$
Definition of loan rate	$(r_{t+1}^b - 1)f^e = f_{t+1}$
Evolution of loan rate	$\hat{r}_{t+1}^b = (1+\phi_r)\hat{r}_t^m$
Households:	
Euler equation (deposit)	$U_{Ct} = \beta E_t(r_{t+1}^m U_{Ct+1})$
Labor supply	$\Sigma_{s=t}^{\infty}(\beta\eta_w)^{s-t}E_t\left\{H_{s\|t}(j)\left[U_{Cs\|t}\frac{W_t^*}{P_s^C} - \chi H_{s\|t}(j)^{1/\phi_l}\right]\right\} = 0$
Market clearing:	
Consumption goods market	$[1-\frac{\eta}{2}(\Phi_t - 1)^2]Y_t = C_t + N_t^e f^e$
Intermediate goods market	$Y_t = X_t$
Loan market	$S_{t+1} = N_t^r f^e$
Aggregate accounting	$C_t + N_t^e v_t + S_{t+1} = w_t L_t + m_t^c - \tau_c Y_t - N_t f_t + r_t^m S_t$

Notes: $DV_{1t} \equiv E_t[\Sigma_{s=t+1}^{\infty}(\Lambda_{t,s}\lambda^{s-t})] = E_t(\Lambda_{t,t+1}\lambda) + E_t(\Lambda_{t,t+1}\lambda DV_{1t+1})$.
$DV_{2t} \equiv E_t[\Sigma_{s=t+1}^{\infty}(\Lambda_{t,s}\lambda^{s-t}f_s)] = E_t(\Lambda_{t,t+1}\lambda f_{t+1}) + E_t(\Lambda_{t,t+1}\lambda DV_{2t+1})$.
$\Gamma_t(z) \equiv \eta\left\{\Phi_t(z)[\Phi_t(z) - 1] - \beta E_t\left[\frac{U_{Ct+1}}{U_{Ct}}\frac{Y_{t+1}}{Y_t}\left(\Phi_{t+1}^C\right)^{\gamma-1}\Phi_{t+1}(z)^{2-\gamma}\left(\Phi_{t+1}(z) - 1\right)\right]\right\}$.
$U_{Ct} \equiv (C_t - bC_{t-1})^{-1} - \beta b[E_t(C_{t+1}) - bC_t]^{-1}$ is the marginal utility of consumption.

3.3. Model solution

3.3.1. Log-linearization

We linearize the model in Table 3.2 by the method of Uhlig (1999). The result is summarized in Table 3.3, where $\pi_t \equiv \frac{p_t - p_{t-1}}{p_{t-1}}$ is the CPI inflation rate, a bar represents a steady-state value. Due to the symmetry assumption in the consumption goods production sector, individual producer price inflation is equal to the average producer price inflation. Therefore, we omit the index z in the notation. We omit z in the notation of other variables in the consumption goods production sector for the same reason. The labor supply equation in the nonlinear model is substituted by two equations. One is the definition of nominal wage inflation π_t^w. The other equation captures the wage inflation dynamics.[42]

[42]See Galì (2008) for a derivation. The only difference is that consumption utility in our model involves habit formation and wage markup is driven to one by the subsidy.

Table 3.3: Log-Linear Model Summary

Consumption goods sector:	
Pricing	$\hat{q}_t = -\hat{\mu}_t$
Markup	$\pi_t = \beta E_t(\pi_{t+1}) - \frac{\gamma}{\eta}\hat{\mu}_t$
Real profit	$\hat{m}^c_t = (\gamma - 1)\hat{\mu}_t + \hat{Y}_t$
Intermediate goods sector:	
Average individual productivity	$\hat{\omega}^m_t = \hat{\omega}^*_t$
Real profit	$\hat{m}_t(\omega^m_t) = \epsilon\hat{q}_t + (1-\epsilon)(\hat{w}_t - \hat{\omega}^m_t - \hat{Z}_t) + \hat{X}_t$
	$\hat{m}_t(\omega^*_t) = \epsilon\hat{q}_t + (1-\epsilon)(\hat{w}_t - \hat{\omega}^*_t - \hat{Z}_t) + \hat{X}_t$
Aggregate pricing	$\hat{q}_t = \frac{1}{1-\epsilon}\hat{N}_t + \hat{w}_t - \hat{\omega}^m_t - \hat{Z}_t$
Firm value	$\hat{U}_{Ct} + \hat{v}_t = E_t(\hat{U}_{Ct+1}) + E_t(\hat{\theta}_{t+1}) + \beta\bar{\theta}\frac{\bar{m}(\bar{\omega}^m)}{\bar{v}}E_t[\hat{m}_{t+1}(\omega^m_{t+1})]$
	$-\beta\bar{\theta}\frac{\bar{f}}{\bar{v}}\hat{f}_{t+1} + \beta\bar{\theta}E_t(\hat{v}_{t+1})$
Free entry	$\hat{v}_t = \hat{w}_t - \hat{Z}_t$
Cutoff condition	$\hat{m}_t(\omega^*_t) = \hat{f}_t$
Survival rate	$\hat{\theta}_t = -k\hat{\omega}^*_t$
Number of firms	$\hat{N}_t = \hat{\theta}_t + \bar{\theta}\hat{N}_{t-1} + (1-\bar{\theta})\hat{N}^e_{t-1}$
Financial sector:	
Fundamental value(gamble)	$\widehat{fv}_t = \widehat{DV}_{1t}$
Bubble size	$(f^g + \overline{DV}_2)\hat{b}b_t + \overline{DV}_2\widehat{DV}_{2t} = 0$
Asset price(gamble)	$\hat{p}^r_t = \hat{b}b_t + \widehat{fv}_t$
Gambling asset market clearing	$f^e\hat{N}^r_t - f^e\bar{\phi}\hat{N}^e_t = f^e(1-\bar{\phi})\hat{P}^r_t$
Proportion of entrepreneur	$\hat{\phi}_t = \hat{N}^e_t - \hat{N}^r_t$
Definition of loan rate	$\bar{r}^b f^e \hat{r}^b_{t+1} = \bar{f}\hat{f}_{t+1}$
Evolution of loan rate	$\hat{r}^b_{t+1} = (1+\phi_r)\hat{r}^m_t$
Households:	
Euler equation (deposit)	$\hat{U}_{Ct} = E_t(\hat{U}_{Ct+1}) + E_t(\hat{r}^m_{t+1})$
Labor supply	$\pi^w_t - \pi_t = \hat{w}_t - \hat{w}_{t-1}$
	$\pi^w_t = \beta E_t\pi^w_{t+1} - \lambda_w(\hat{w}_t - \frac{1}{\phi_l}\hat{L}_t + \hat{U}_{Ct})$
Market clearing:	
Consumption goods market	$\bar{Y}\hat{Y}_t = \bar{C}\hat{C}_t + \bar{N}^e f^e \hat{N}^e_t$
Intermediate goods market	$\hat{X}_t = \hat{Y}_t$
Loan market	$\hat{S}_{t+1} = \hat{N}^r_t$
Aggregate accounting	$\bar{C}\hat{C}_t + \bar{N}^e\bar{v}(\hat{N}^e_t + \hat{v}_t) + \bar{S}\hat{S}_{t+1} = \bar{w}\bar{L}(\hat{w}_t + \hat{L}_t)$
	$+\bar{m}^c(\hat{m}^c_t - \hat{Y}_t) - \bar{N}\bar{f}(\hat{N}_t + \hat{f}_t) + \bar{r}^m\bar{S}(\hat{r}^m_t + \hat{S}_t)$

Notes: $\widehat{DV}_{1t} = (1-\beta\lambda)E_t(\hat{U}_{Ct+1} - \hat{U}_{Ct}) + \beta\lambda E_t(\hat{U}_{Ct+1} - \hat{U}_{Ct} + \widehat{DV}_{1t+1})$.
$\widehat{DV}_{2t} = (1-\beta\lambda)E_t(\hat{U}_{Ct+1} - \hat{U}_{Ct} + \hat{f}_{t+1}) + \beta\lambda E_t(\hat{U}_{Ct+1} - \hat{U}_{Ct} + \widehat{DV}_{2t+1})$.
$\hat{U}_{Ct} = \frac{\beta b E_t(\hat{C}_{t+1}) - (1+\beta b^2)\hat{C}_t + b\hat{C}_{t-1}}{(1-\beta b)(1-b)}$ is the deviation of log marginal utility of consumption from its steady-state level.
$\lambda_w \equiv \frac{(1-\eta_w)(1-\beta\eta_w)}{\eta_w(1+\epsilon_w/\phi_l)}$.

3.3.2. Calibration

As in the standard business cycle model, the periods are interpreted as quarters. The household discount factor β is set to 1/1.0025, which implies that the US steady-state monetary policy rate is 1% per annum (Goodfriend and McCallum, 2007; Curdia and Woodford, 2010). The habit formation parameter b is set to 0.65, the value estimated by Christiano et al. (2005). We set the elasticities $\gamma = \epsilon = 3.8$ to fit the U.S. plant and macro trade data (Ghironi and Melitz, 2005; Bilbiie et al., 2008). Following Ghironi and Melitz (2005), we calibrate the Pareto distribution shape parameter k to match the standard deviation of log US plant sales which is 1.67 according to Bernard et al. (2003). We follow Bilbiie et al. (2008) by setting the Frisch elasticity to $\phi_l = 2$ and the price stickiness parameter to $\eta = 77$.[43] The weight of labor disutility χ is calibrated to generate a steady-state labor effect level of one regardless of the Frisch elasticity. Following Erceg et al. (2000), we set elasticity of substitution across labor types ϵ_w to 4 and sticky wage parameter η_w to 0.75. The parameter governing interest rate passthrough, ϕ_r , is set to 0.3, the value estimated for the US by Chowdhury et al. (2006). The steady-state lending rate is set to $(1.02)^{1/4}$ times the monetary policy rate, reflecting the 2% US steady-state annual credit spread (Curdia and Woodford, 2010). We normalize the sunk cost in consumption goods f^e to 1 since its level does not affect the coefficients of the impulse response functions. The steady-state intermediate goods producer survival rate is set to 0.975, the same number as the one specified in Ghironi and Melitz (2005) and Bilbiie et al. (2008). The difference is that firms' survival rate is fixed in Ghironi and Melitz (2005) and Bilbiie et al. (2008) while it can deviate from the steady-state level in our model. We require the calibrated steady-state variables to capture the US private debt-to-GDP ratio (S_t/y_t^a), 80% per annum or 3.2 per quarter (Curdia and Woodford, 2010). We set $\lambda = 0.9$ and $f^g = 0.934$, so that there is no bubble in the steady state. The periodic supply of the gambling asset GS is normalized to one since it does not affect the impulse response function.[44] Following King and Rebelo (1999) and Bilbiie et al. (2008), we assume that the aggregate productivity evolves as follows: $lnZ_t = 0.979 lnZ_{t-1} + e$, where e is an i.i.d. random shock with variance σ^2.

Given those specifications of parameters and steady-state variables, we can solve for the sunk cost in labor f^{ew} and other steady-state variables.[45] Table 3.4 summarizes the steady-state magnitudes. The quarterly private debt-to-GDP ratio is $\bar{S}/\bar{y}^a \equiv \frac{\bar{S}}{\bar{c}+\bar{n}^e\bar{v}} \doteq 3.2$, as we required. The consumption-to-GDP ratio is $\bar{c}/\bar{y}^a \equiv \frac{\bar{c}}{\bar{c}+\bar{n}^e\bar{v}} \doteq 0.717$, which roughly equals the US data. The periodic repayment satisfies the definition $\bar{f} = (\bar{r}^b - 1)f^e$. The price markup of the intermediate goods sector equals 1 because the distortion

43 The key results are robust to different values of the Frisch elasticity ranging from 0 to 3.

44 What matters for the impulse responses is the product of steady-state price and periodic supply of the gambling asset. This product is determined by the steady-state gambling asset market clearing condition, given the numerical values of our other parameters.

45 The calibrate value of f^{ew} is 0.04.

from monopolistic competition is eliminated by the production subsidy. It is also easy to verifies that $\bar{m}^c - \tau_c \bar{Y}$ is zero. As we discussed before, this result is due to the fact that the price markup is driven to one by the production subsidy.

Table 3.4: Steady state magnitudes

$\bar{r}^m$	$\bar{r}^b$	$\bar{\theta}$	$\bar{S}$	$\bar{c}$	$\bar{f}$	$\bar{\mu}$	$\bar{w}$
1.0025	1.0075	0.975	25.464	5.757	0.0075	1	9.516
$\bar{m}^c$	$\bar{Y}$	$\bar{v}$	$\bar{\omega}^m$	$\bar{n}$	$\bar{n}^e$	$\bar{n}^r$	
4.043	11.32	0.396	1.393	216.984	5.564	25.464	

3.4. Impulse responses

We consider the impulse responses of the variables to two types of shocks: (1) an expansionary monetary policy shock; and (2) a negative productivity shock. We focus on the variables related to the riskiness of loan portfolios for financial intermediaries. More specifically, we report and discuss the impulse responses of the number of entrepreneurs entering, the number of entering gamblers, the survival rate of entrepreneurs, the bubble size, and the gambling asset price. We also report the impulse responses of two other variables (the average profit of intermediate goods producers and the required periodic repayment to financial intermediaries) since they are closely related to investors' entry decisions. [46]

3.4.1. An expansionary monetary policy shock

Figure 3.2 shows the impulse responses of the key variables (percentage deviations from the steady-state levels) to a one percent unexpected decrease in the net nominal money market interest rate.[47][48] The variable on the horizontal axis is the number of years after the shock. The responses are normalized so that one denotes one percent. The dashed

[46] Interested readers can refer to the figures in the appendices for the impulse responses of all the other variables in the model.

[47] Note that here we are talking about a one percent shock to the net interest rate rather than the gross one. A one percent shock to the gross interest rate will reduce the quarterly net nominal interest rate to a negative value, which is not possible in reality. To see this, notice that the steady state gross nominal rate is 1.0025. Reducing this number by one percent gives a nominal gross rate of 0.9925.

[48] See figure A1 for impulse responses of all variables.

curves with square markers correspond to the responses to the shocks under the interest rate smoothing rule without reacting to output fluctuations. The solid curves with cross markers correspond to the responses under the interest smoothing rule reacting to output fluctuations. The dotted curves with round markers correspond to the responses under the forward-looking Taylor rule.

The impulse responses are qualitatively similar under all three interest rate rules. More specifically, after the shock, entrepreneurial entry increases in the first few quarters, and then starts to decline and persistently stays below the steady-state level for a long period of time. This result is in sharp contrast to the result of Bilbiie et al. (2008). Bilbiie et al. (2008) who find that the expansionary monetary policy shock immediately reduces firm entry if entry incurs sunk investment in effective labor. Furthermore, they find firm entry persistently stays above the steady-state level after the first few periods. As noted by Rotemberg (2008), the initial decline in firm entry in Bilbiie et al. (2008) comes from the procyclical rise in the real wage, which makes entry more expensive and future returns less attractive. Particularly, average profit in the intermediate goods sector decreases after the shock despite the increase in demand. If the nominal wage is sticky, the real wage is less procyclical and entry is less costly. Additionally, average profits in the near future rise. This further encourages entry. Our result confirms Rotemberg's conjecture that an expansionary monetary policy shock stimulates entry on impact when a realistic level of wage rigidity is introduced into the model. Holding the number of producers constant, the increase in demand also increases profitability in the current period. This makes debt repayment easier and raises the survival rate of intermediate goods firms. Both the increase in firm survival rate and entrepreneurial entry in the initial period increase the number of producers in the future. Intensified competition not only reduces the sales of each individual intermediate goods producer but also reduces the price of intermediate goods relative to consumption goods. This is because a one percent decrease in the price of aggregate intermediate goods only leads to a less than one percent decrease in the price of aggregate consumption goods when prices in the consumption goods sector are sticky. As a result, future profits in real terms decrease, leading to lower levels of entrepreneurial entry.

The number of gamblers in the pool of new loan applicants is persistently higher than the steady-state level after the shock. This is because the unexpected decline in the nominal interest rate persistently reduces required periodic repayments to the financial intermediaries.[49] Lower periodic repayments lead to higher expected cash flows from the gamble, which attracts more gamblers. The increase in the number of gamblers pushes up the price of the gambling asset. Limited liability encourages excessive risk-taking behavior. Therefore, the rise in the gambling asset price is more than the rise in its fundamental value. In other words, the bubble size becomes larger than one. Recall that the expected repayment from gamblers to the financial intermediaries is inversely related to the bubble size. Therefore, after the expansionary monetary policy shock, the expected loss from lending to gamblers will be persistently higher than the steady-state level. Figure 3.2 suggests that the initial rise in the number of entrepreneurs quantitatively dominates the initial rise in the number of gamblers. Therefore, the proportion of entrepreneurs in the borrower pool initially increases. However, the initial increase in the proportion of entrepreneurs does not last long. Instead, the proportion of entrepreneurs persistently stays below the steady-state level in the long run. A persistently higher than steady state expected loss from lending to gamblers together with a persistently higher than steady state proportion of gamblers in the borrower pool accumulates a significant risk in the financial sector. Interestingly, the effect of the monetary policy shock on the accumulation of long-run financial risk is quantitatively much more significant when the interest rate does not react to output fluctuations. This is because under the rules reacting to the output fluctuation, initial rise in entrepreneurial entry is reduced by the central bank's action to cut aggregate demand. The lower initial rise in entrepreneurial entry reduces the future numbers of competitors in the market, making entry in the following periods more attractive. Taylor (2009) argues that keeping the policy interest rate persistently lower than the level implied by the Taylor rule may be a source of financial crisis. We find that if the economy is hit by an expansionary monetary policy shock and the central bank does not react to output fluctuations, the nominal money market interest rate will be persistently lower than the level implied by the forward-looking Taylor rule.[50] As we discussed, not reacting to output fluctuations leads to a more significant long-run financial risk. In this sense, our findings in this section are consistent with

[49] Required periodic repayments increase in the initial periods. However, the effect of the persistent reduction in future required periodic repayments dominates the changes in the net present value of cash flows.

[50] See figure A1.

Taylor's argument. However, we shall see in the next subsection, sticking to the Taylor rule is not sufficient to eliminate financial crises.

3.4.2. A negative productivity shock

Figure 3.3 shows the impulse responses of the key variables (percentage deviations from the steady-state levels) to a one standard deviation[51] decrease in aggregate productivity, Z_t.[52] The first observation is that impulse responses are very similar under the two interest rate rules reacting to output fluctuations. Second, impulse responses under the two rules reacting to output fluctuations are very different from the ones under the the interest rate rule not reacting to output fluctuations. More specifically, we find the following key results.

When the interest rate rule reacts to output fluctuations, entrepreneurial entry initially increases. By contrast, entrepreneurial entry initially decreases when the interest rate rule does not react to output fluctuations. The initial decrease in aggregate productivity affects entrepreneurial entry through two channels. The first one is the ***direct profit channel***: persistently lower than steady state aggregate productivity can directly reduce future profitability of intermediate goods production, which deters entrepreneurial entry. The second channel is the ***interest rate channel***: the real money market interest rate decreases after the shock under all three interest rate rules. The reduction in the real money market rate reduces future real loan rates and required periodic repayments, making entrepreneurial entry more attractive. Additionally, lower real money market rates increase demand. This reduces the negative effect of the productivity shock on production and profits and further encourages entrepreneurial entry. The net effect of the negative aggregate productivity shock on entrepreneurial entry depends on the size of the offsetting effects. If the interest rate reacts to output fluctuations, the interest rate channel dominates on impact and the firm value exceeds the sunk cost of investment, which means that entrepreneurial entry must increase to preserve the free entry condition in the intermediate goods sector. By contrast, if the interest rate rule does not react to output fluctuations, the direct profit channel dominates on impact, leading to an immediate reduction in entrepreneurial entry.

[51]The standard deviation of the aggregate productivity shock is set to 0.0012, the number used in King and Rebelo (1999).

[52]See figure A2 for impulse responses of all variables.

The initial increase in entrepreneurial entry under the two interest rate rules reacting to output fluctuations do not last long and are followed by persistently lower than steady state numbers of entrepreneurial entry. This is because the initial rise in entrepreneurial entry makes the number of intermediate goods producers persistently higher than the steady-state number. Competition reduces future profitability and deters entry. By contrast, under the interest rate rule not reacting to output fluctuation, due to the initial decrease in entrepreneurial entry, the number of intermediate goods producers is persistently below its steady-state value. Less competition attracts entry, so entrepreneurial entry quickly recovers and remains at higher than steady state values for a long period of time.

The firm survival rate initially increases after the negative aggregate productivity shock under the interest rate rules reacting to output fluctuations, whereas it initially decreases under the interest rate rule not reacting to output fluctuations. The responses of the firm survival rate become quantitatively very small after five years under all interest rate rules. Similar to entrepreneurial entry, the firm survival rate is also affected by the bad productivity shock through two channels: the direct profit channel and the interest rate channel. Lower productivity reduces profits while the lower interest rate increases profits by increasing demand.[53] If the interest rate rule reacts to output fluctuations, the interest rate channel initially dominates, leading to a higher firm survival rate. Conversely, if the interest rate rule does not react to output fluctuations, the direct profit channel dominates on impact. As a result, the firm survival rate decreases. A higher survival rate of the firms also increases the future number of competitors and deters entrepreneurial entry in the long run.

The number of gamblers initially increases under all three different monetary policy rules. However, the initial rise in the number of gamblers is small and transitory if the central bank does not react to output fluctuations. By contrast, the initial rise in the number of gamblers is large and persistent if the central bank does react to output fluctuations. Consequently, the bubble size is persistently higher if the central bank reacts to output fluctuations. The intuition is as follows. Reduction in real interest rates reduces future required repayments and increases cash flows from gambling. This attracts

[53]Note that required repayment is predetermined when the shock hits, so the cut in real interest rate does not work through affecting the required repayment in the initial period.

gamblers. Excessive risk-taking behavior by the gamblers increases the bubble size. If the central bank tries to avoid the current recession by cutting the nominal interest rate, it lowers the real interest rate more than when it does not care about output fluctuations. As a result, cash flows from gambling increase more and more gamblers enter the market, leading to a larger size of the bubble. A larger bubble size suggests a higher expected loss from lending to gamblers. Together with a higher proportion of gamblers in the borrower pool, it imposes a significant risk to the financial sector. The results suggest that sticking to a Taylor rule is not sufficient to eliminate financial crises. Actually, in case the economy is hit by a negative productivity shock, deviating from the Taylor rule by not reacting to output fluctuations can reduce the long-run financial risk.

3.5. Sticky interest rate pass-through

In our benchmark model, the pass-through from changes in money market rate to the loan rate is more than one. It is interesting to see what happens if we have a lower interest rate pass-through. Particularly, many studies find that the interest rate pass-through is sticky in Europe. In this section, we ask the question: "other things being equal (we keep other parameter values calibrated for the US economy), what is the impact of interest rate pass-through on the impulse responses?" More specifically, we produce impulse responses of the variables to the shocks with $\phi_r = -0.8$ which implies an interest rate pass-through of 0.2, the value estimated by Chowdhury et al. (2006) for France. [54]

3.5.1. An expansionary monetary policy shock

Figure 3.4 displays the impulse responses of key variables after a one percent unexpected decrease in net nominal money market rate.[55] The qualitative results are similar to the benchmark model. The number of entrepreneurs entering initially rises and then remains at levels lower than the steady-state value for a long time. The intermediate firm survival rate initially rises, followed by quantitatively negligible responses. The

[54] An interest rate pass-through of 0.2 is the lower bound found in the sample of Chowdhury et al. (2006). Hence, the results of this section show that our key findings in the previous section are robust to a wide range (from 0.2 to 1.3) of values of the interest rate pass-through.

[55] See figure A3 for impulse responses of all variables.

number of gamblers in the borrower pool increases and stays at levels higher than the steady-state level for a long time. The proportion of entrepreneurs initially increases, starts to decline after a short period and remains at levels lower than the steady-state level for a long time. Bubble size and real asset price increases, and persistently stay at levels higher than the steady-state levels.

Two notable differences from the benchmark model are: variables converge to their steady-state levels faster than in the benchmark model; the quantitative responses are less different under the three interest rate rules than in the benchmark model. This is because now the differences in the effects of initial change in money market rate are narrowed down by the sticky interest passthrough when transmitted to the intermediate goods sector.

3.5.2. A negative productivity shock

Figure 3.5 displays the impulse responses of key variables after a one standard deviation negative productivity shock.[56] As in the benchmark model, entrepreneurial entry initially increases, then declines to a level lower than the steady-state level, and slowly recovers when the nominal interest rate reacts to output fluctuations. The difference is that the initial increase in entrepreneurial entry is smaller, leading to smaller numbers of future competitors in the intermediate goods sector. Hence, the proportion of entrepreneurs in the borrower pool converges faster to the steady-state level than in the benchmark model. The bubble size remains above the steady state for more than five years if the interest rate reacts to output fluctuations. However, both size and duration of the bubble are smaller in magnitude than in the benchmark model. Therefore, when the interest rate passthrough is sticky, the economy hit by a negative productivity shock is less prone to a long-run financial crash.

3.6. Conclusion

Our model demonstrates that large unexpected expansionary monetary policy shocks could trigger financial crises in the long run. Interestingly, the central bank's reaction to output fluctuations can reduce the negative effect of the unexpected reduction in money

[56] See figure A4 for impulse responses of all variables.

market interest rate on the long-run financial stability. As we know, the Taylor rule includes the central bank's reaction to output fluctuations. In this sense, sticking to the Taylor rule can help reduce the long-run financial risk. However, a central bank's monetary policy aimed at smoothing output fluctuations can persistently worsen the borrower pool faced by financial intermediaries in the long run if the economy is hit by a negative aggregate productivity shock. That is, it will persistently increase the proportion of gamblers in the pool of new loan applicants. Furthermore, the expected loss from lending to each gambler is persistently higher than the steady-state level under such a policy. If the central bank only responds to inflation, the negative effect of the aggregate productivity shock on the borrower pool is more transitory but larger in magnitude, which suggests that the financial intermediaries have to temporarily withstand higher pressure. As a tradeoff, they can avoid persistent future losses if they survive the current stress. The traditional business cycle view of financial crises[57] suggests that a sharp drop in the productivity of the real sector could generate a bank run. Hence, it is tempting for governments to intervene to avoid financial crises. However, our analysis suggests that policies that try to reduce the probability of a current crisis may create a future crisis in the long run.

[57] See Allen and Gale (2007) for a summary.

Figure 3.1: Structure of the model

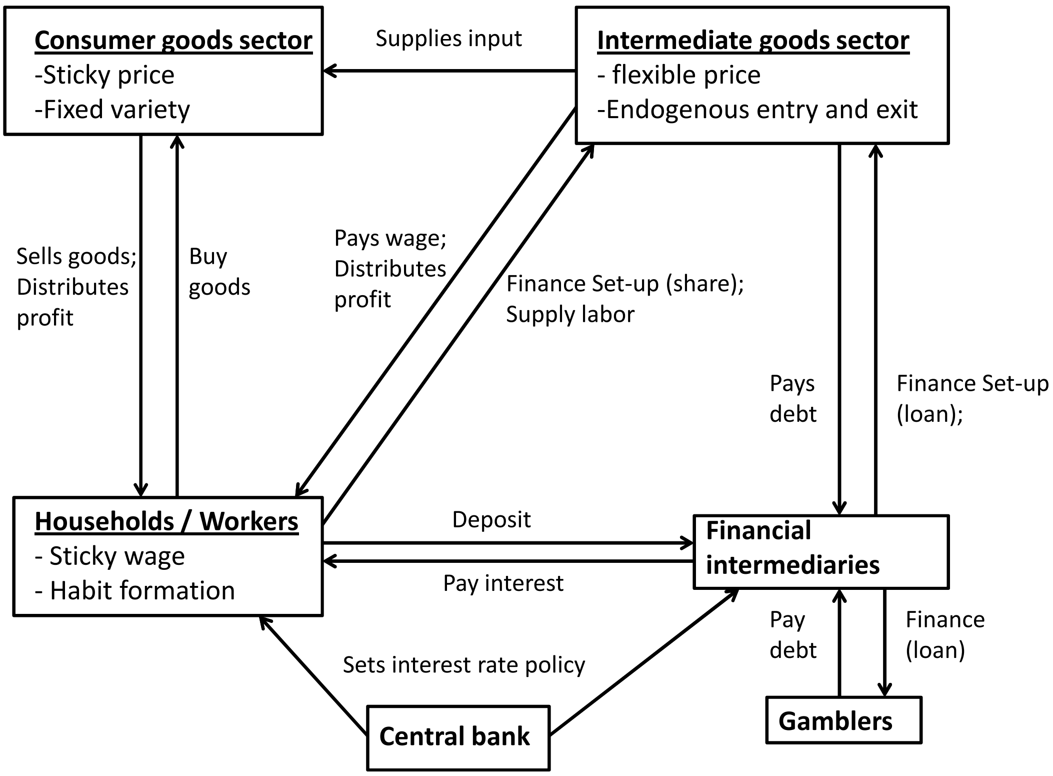
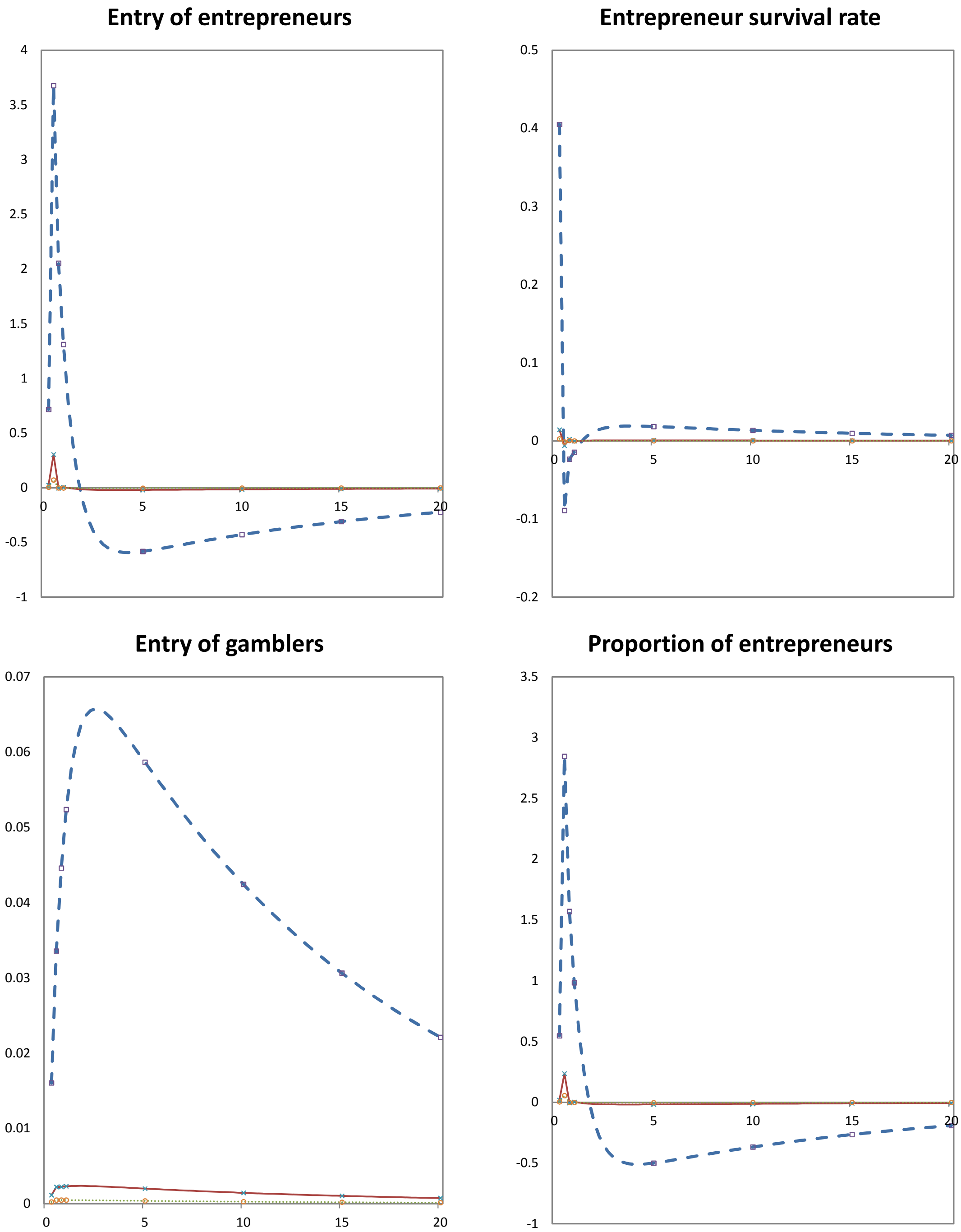
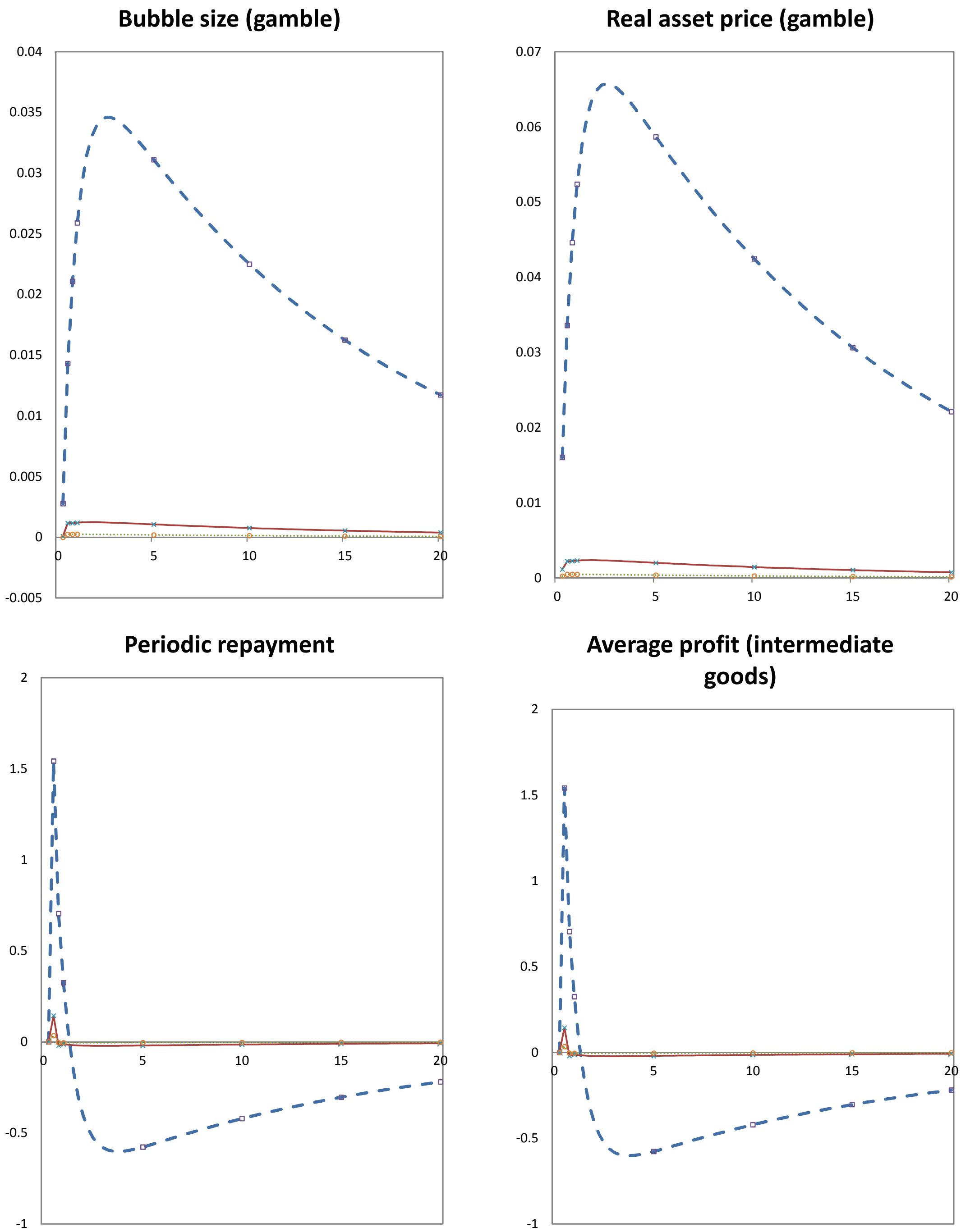


Figure 3.2: Impulse Responses After An Expansionary Monetary Policy Shock, Key Variables



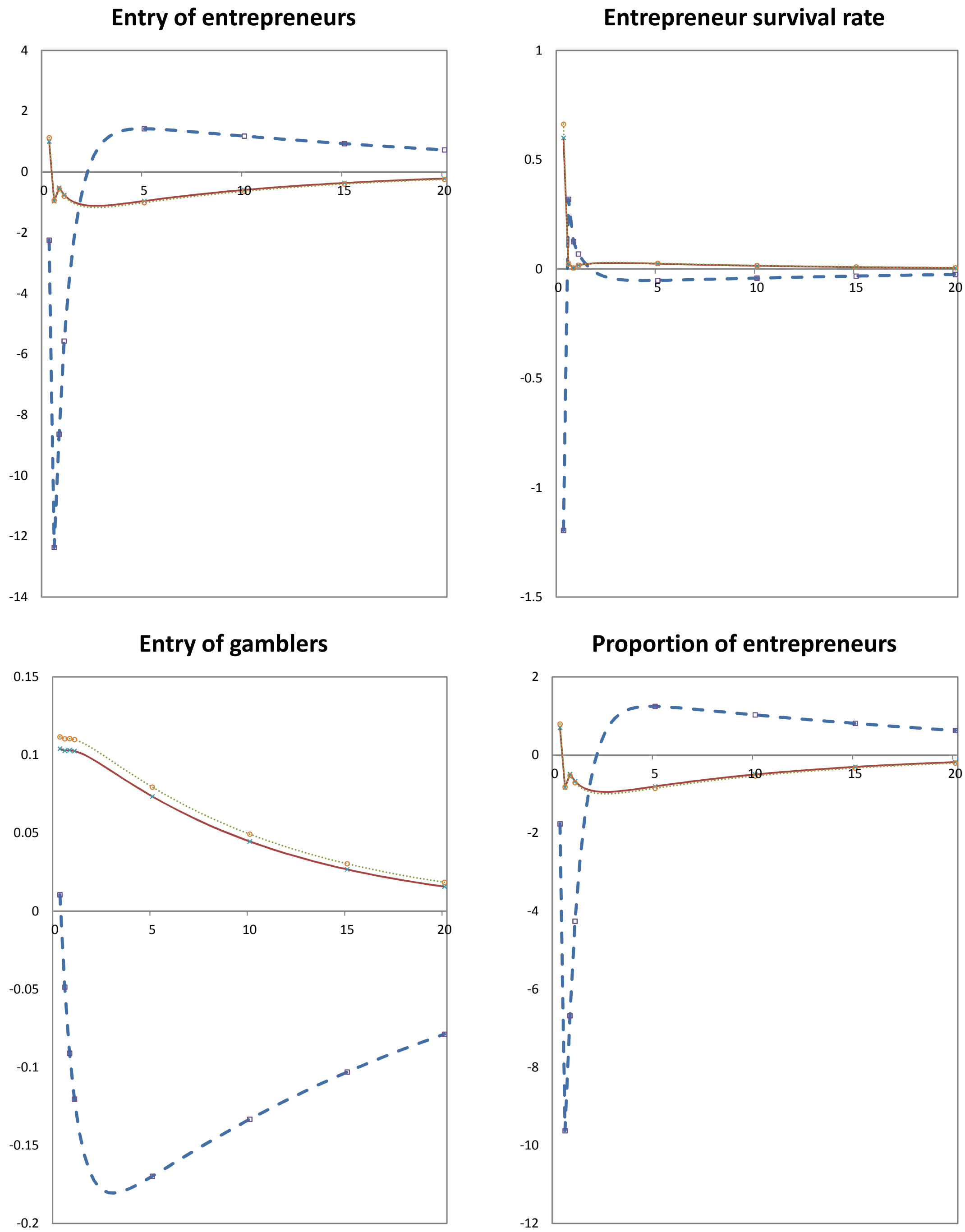
Notes: The variable on the horizontal axis is the number of years after the shock. The responses are normalized so that one denotes one percent deviation from the steady-state level. The dashed curves with square markers correspond to the responses to the shocks under the interest rate smoothing rule without reacting to output fluctuations. The solid curves with cross markers correspond to the responses under the interest smoothing rule reacting to output fluctuations. The dotted curves with round markers correspond to the responses under the forward-looking Taylor rule.

Impulse Responses After An Expansionary Monetary Policy Shock, Key Variables (Continued)



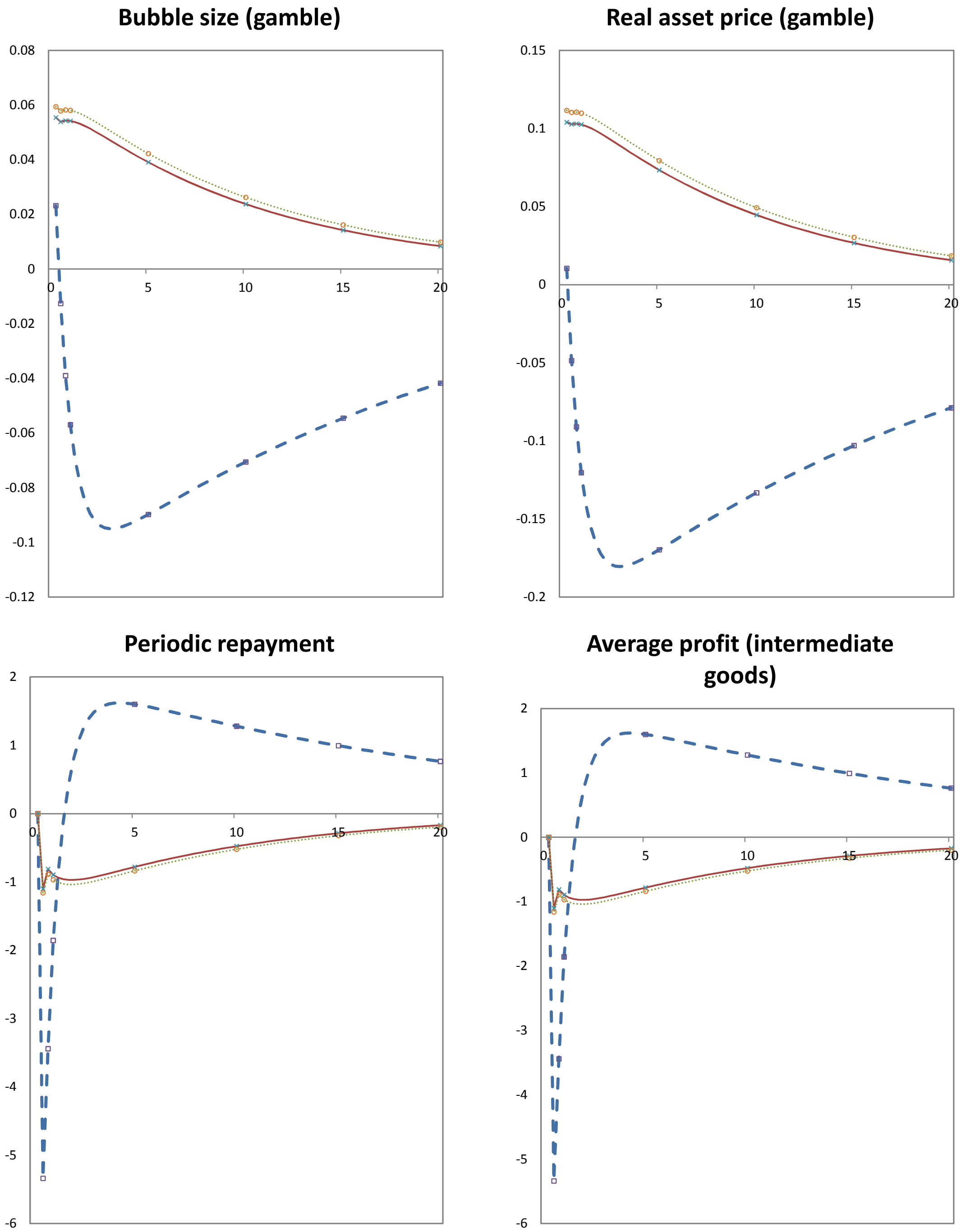
Notes: The variable on the horizontal axis is the number of years after the shock. The responses are normalized so that one denotes one percent deviation from the steady-state level. The dashed curves with square markers correspond to the responses to the shocks under the interest rate smoothing rule without reacting to output fluctuations. The solid curves with cross markers correspond to the responses under the interest smoothing rule reacting to output fluctuations. The dotted curves with round markers correspond to the responses under the forward-looking Taylor rule.

Figure 3.3: Impulse Responses After A Negative Productivity Shock, Key Variables



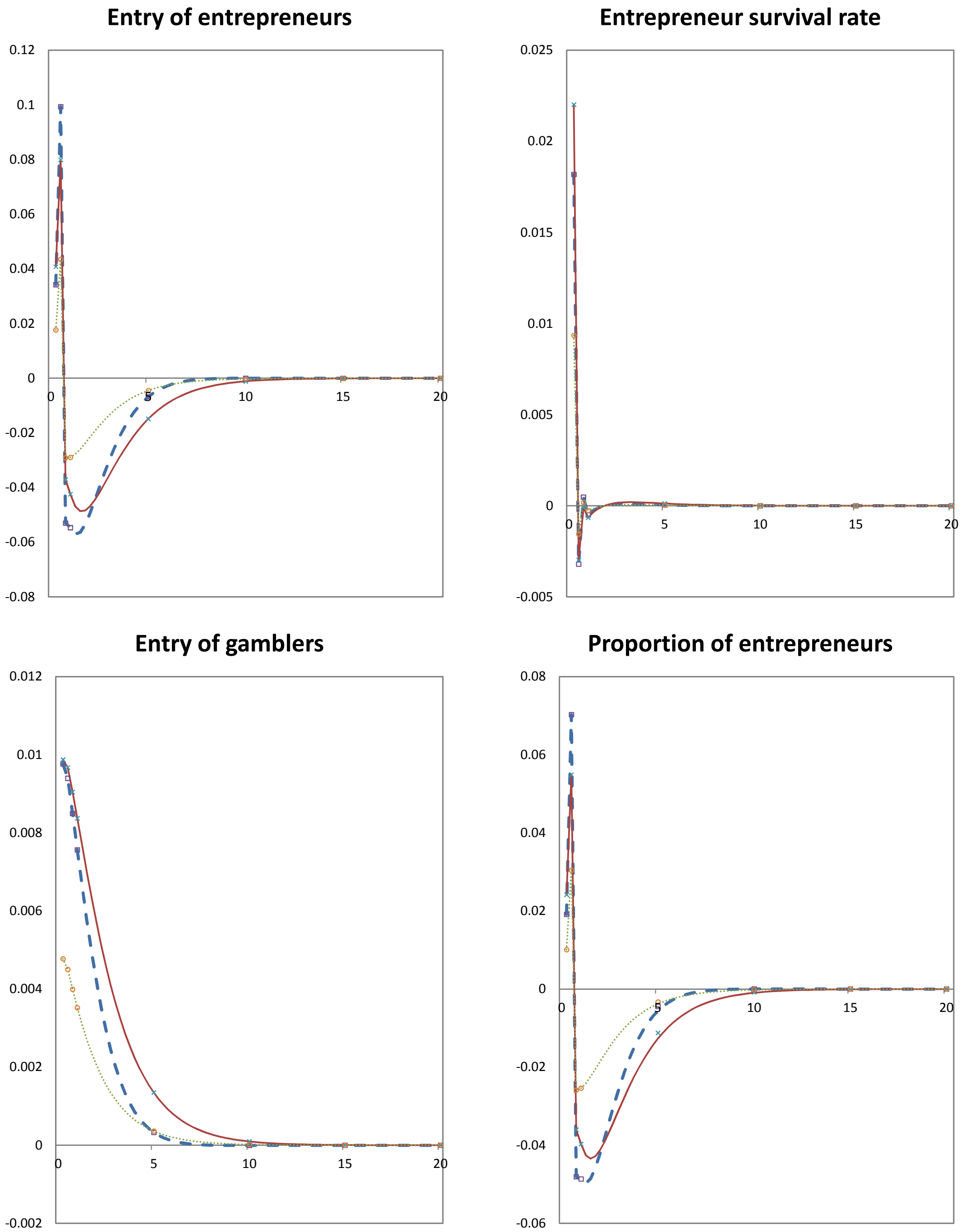
Notes: The variable on the horizontal axis is the number of years after the shock. The responses are normalized so that one denotes one percent deviation from the steady-state level. The dashed curves with square markers correspond to the responses to the shocks under the interest rate smoothing rule without reacting to output fluctuations. The solid curves with cross markers correspond to the responses under the interest smoothing rule reacting to output fluctuations. The dotted curves with round markers correspond to the responses under the forward-looking Taylor rule.

Impulse Responses After A Negative Productivity Shock, Key Variables (Continued)



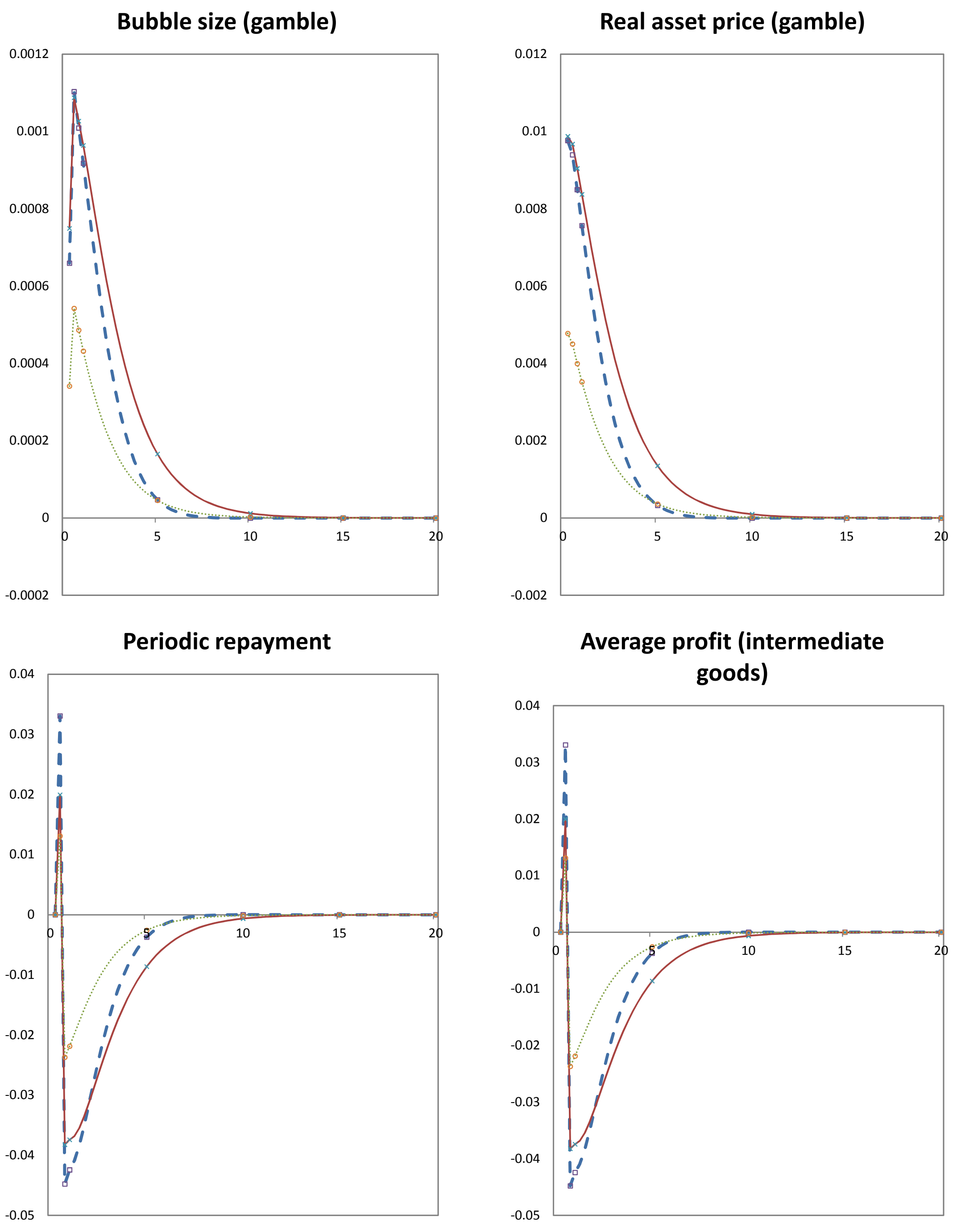
Notes: The variable on the horizontal axis is the number of years after the shock. The responses are normalized so that one denotes one percent deviation from the steady-state level. The dashed curves with square markers correspond to the responses to the shocks under the interest rate smoothing rule without reacting to output fluctuations. The solid curves with cross markers correspond to the responses under the interest smoothing rule reacting to output fluctuations. The dotted curves with round markers correspond to the responses under the forward-looking Taylor rule.

Figure 3.4: Sticky Interest Rate Passthrough and Impulse Responses After An Expansionary Monetary Policy Shock, Key Variables



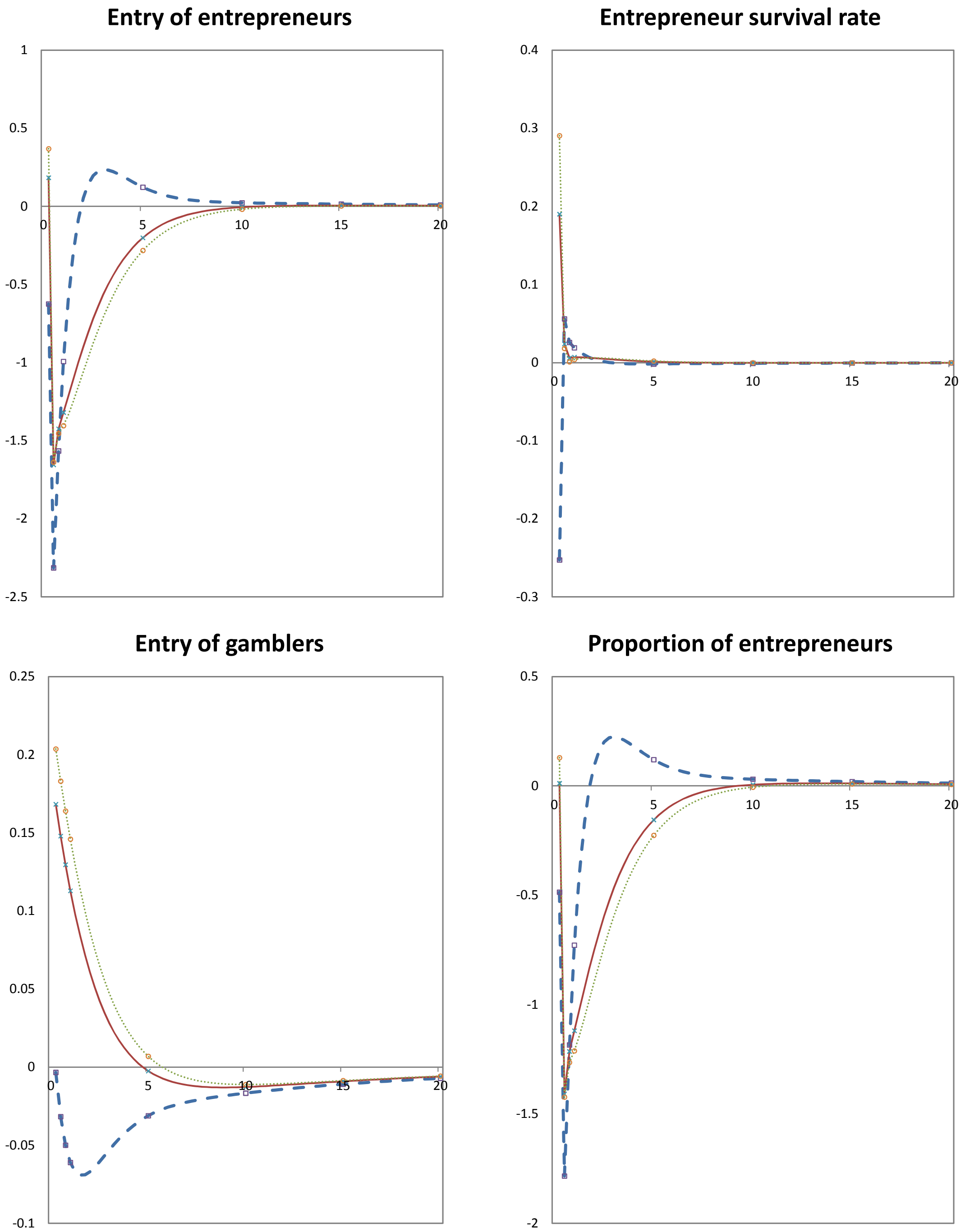
Notes: The variable on the horizontal axis is the number of years after the shock. The responses are normalized so that one denotes one percent deviation from the steady-state level. The dashed curves with square markers correspond to the responses to the shocks under the interest rate smoothing rule without reacting to output fluctuations. The solid curves with cross markers correspond to the responses under the interest smoothing rule reacting to output fluctuations. The dotted curves with round markers correspond to the responses under the forward-looking Taylor rule.

Sticky Interest Rate Passthrough and Impulse Responses After An Expansionary Monetary Policy Shock, Key Variables (Continued)



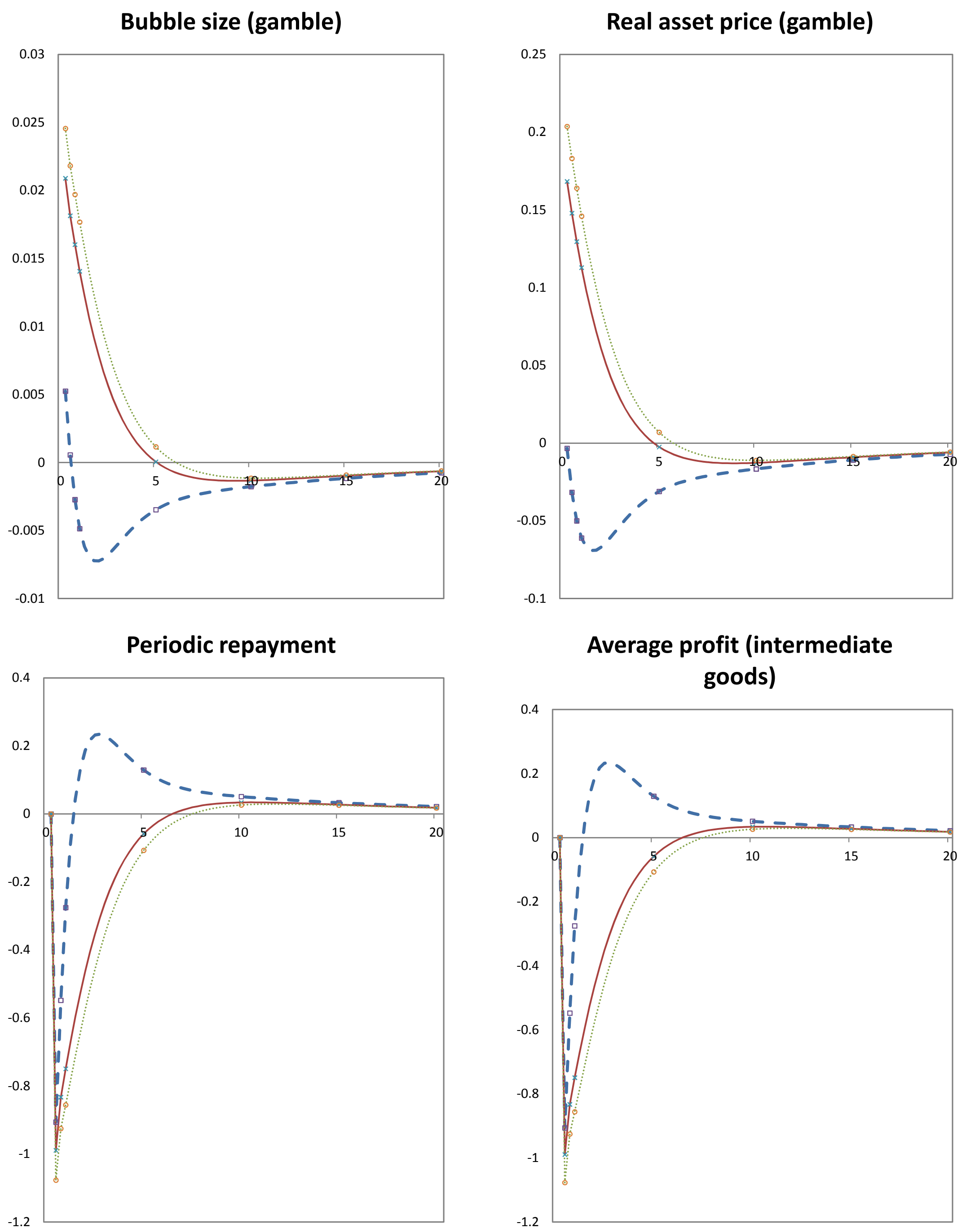
Notes: The variable on the horizontal axis is the number of years after the shock. The responses are normalized so that one denotes one percent deviation from the steady-state level. The dashed curves with square markers correspond to the responses to the shocks under the interest rate smoothing rule without reacting to output fluctuations. The solid curves with cross markers correspond to the responses under the interest smoothing rule reacting to output fluctuations. The dotted curves with round markers correspond to the responses under the forward-looking Taylor rule.

Figure 3.5: Sticky Interest Rate Passthrough and Impulse Responses After A Negative Productivity Shock, Key Variables



Notes: The variable on the horizontal axis is the number of years after the shock. The responses are normalized so that one denotes one percent deviation from the steady-state level. The dashed curves with square markers correspond to the responses to the shocks under the interest rate smoothing rule without reacting to output fluctuations. The solid curves with cross markers correspond to the responses under the interest smoothing rule reacting to output fluctuations. The dotted curves with round markers correspond to the responses under the forward-looking Taylor rule.

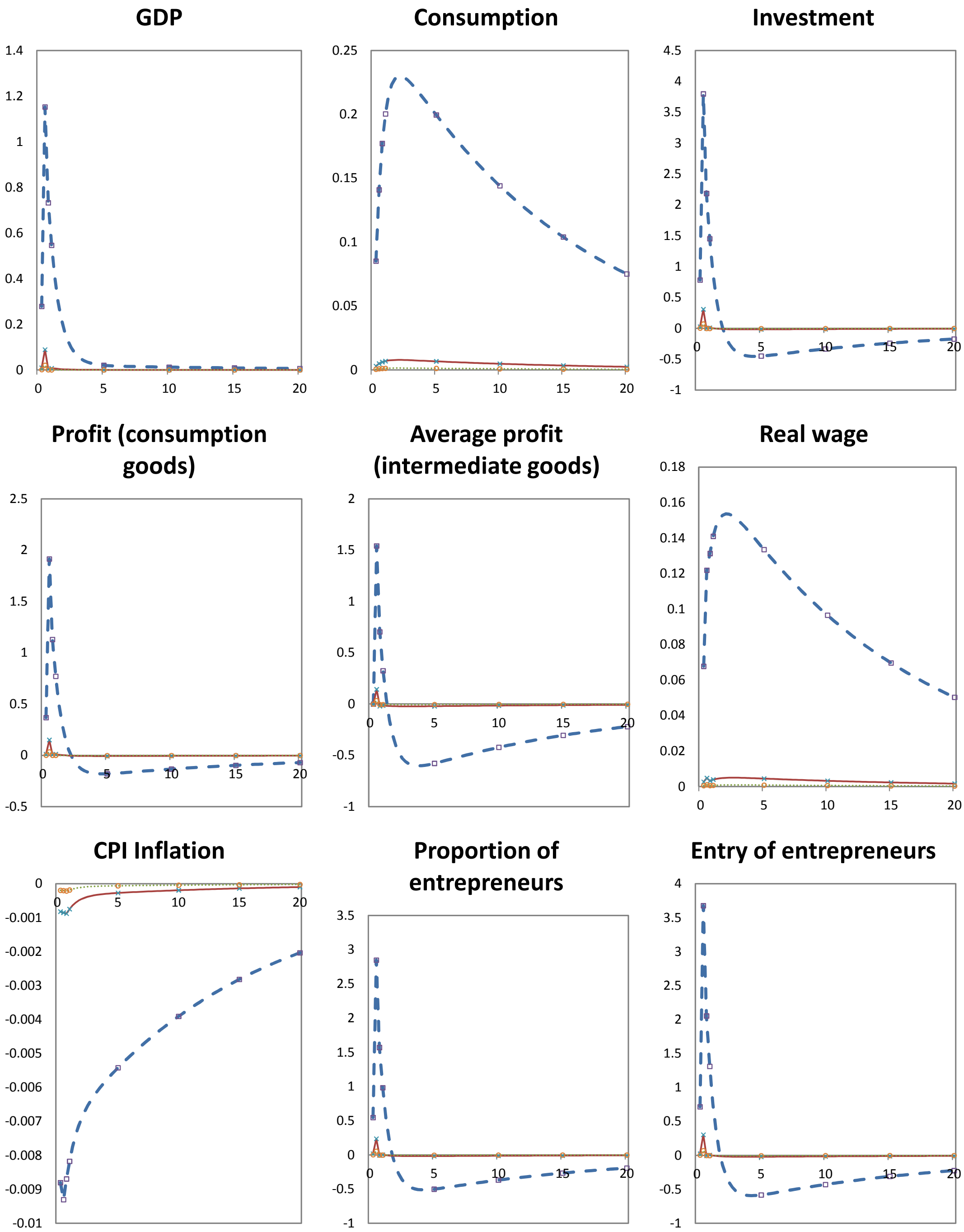
Sticky Interest Rate Passthrough and Impulse Responses After A Negative Productivity Shock, Key Variables (Continued)



Notes: The variable on the horizontal axis is the number of years after the shock. The responses are normalized so that one denotes one percent deviation from the steady-state level. The dashed curves with square markers correspond to the responses to the shocks under the interest rate smoothing rule without reacting to output fluctuations. The solid curves with cross markers correspond to the responses under the interest smoothing rule reacting to output fluctuations. The dotted curves with round markers correspond to the responses under the forward-looking Taylor rule.

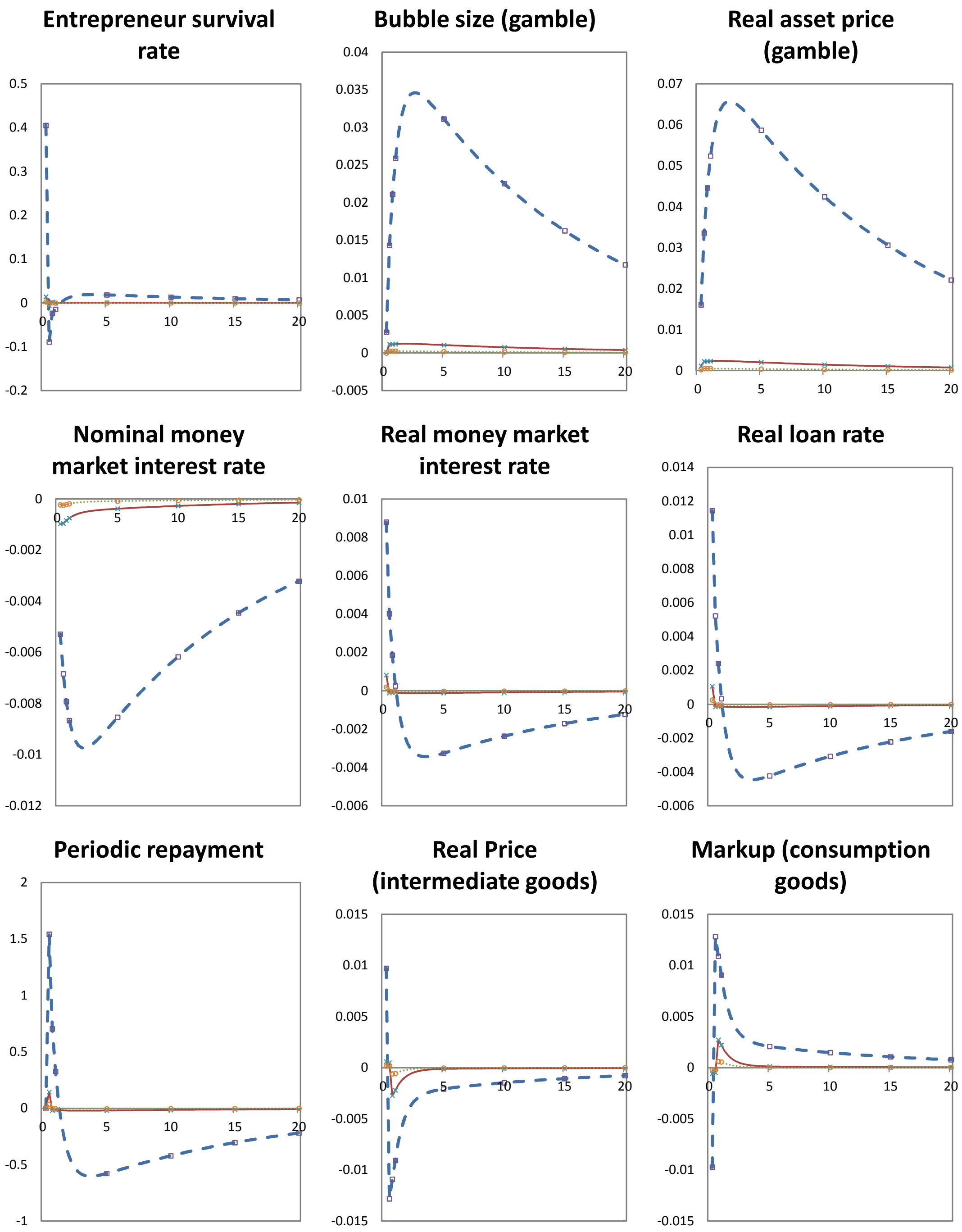
3.7. Appendix

Figure A1: Impulse Responses After An Expansionary Monetary Policy Shock



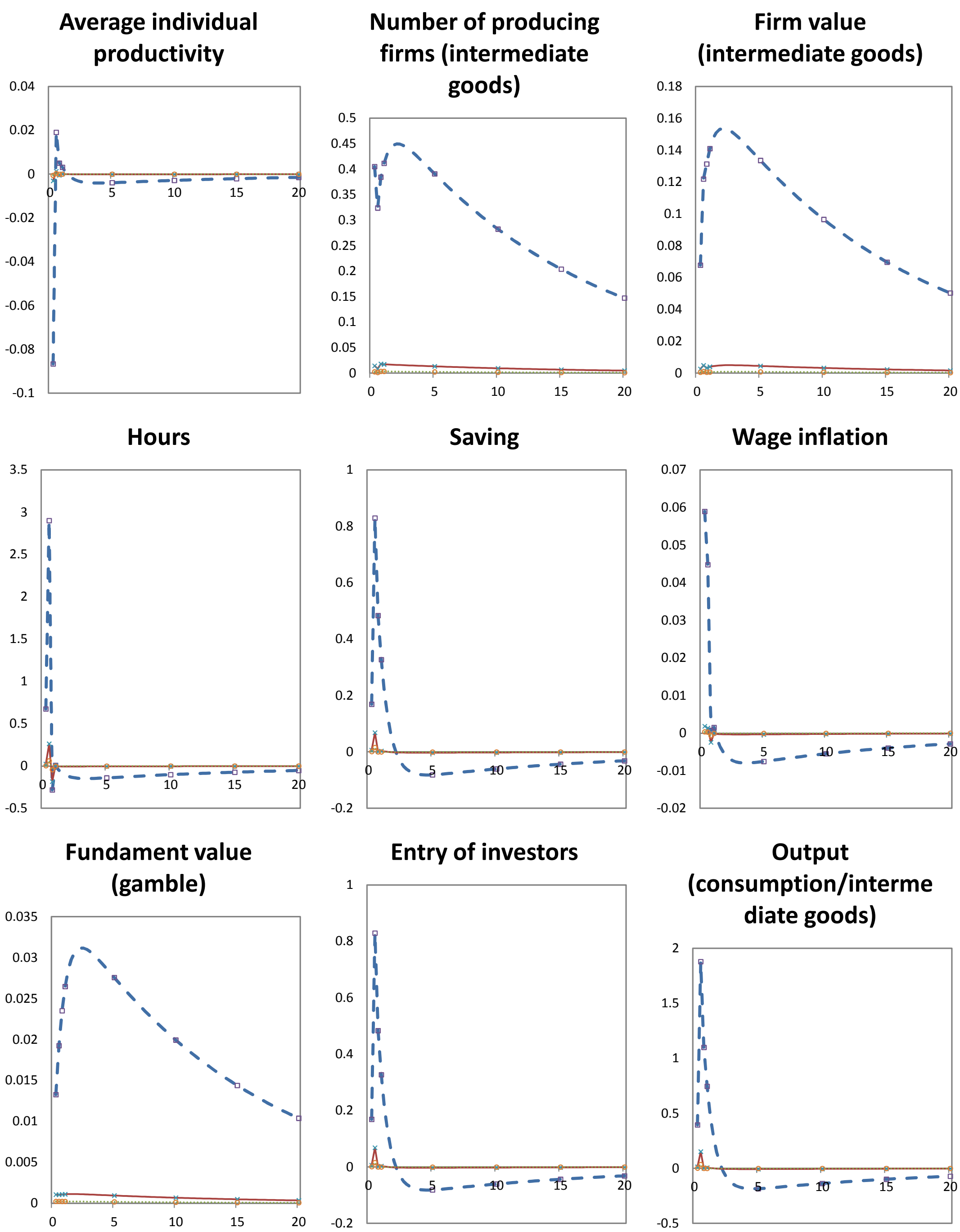
Notes: The variable on the horizontal axis is the number of years after the shock. The responses are normalized so that one denotes one percent deviation from the steady-state level. The dashed curves with square markers correspond to the responses to the shocks under the interest rate smoothing rule without reacting to output fluctuations. The solid curves with cross markers correspond to the responses under the interest smoothing rule reacting to output fluctuations. The dotted curves with round markers correspond to the responses under the forward-looking Taylor rule.

Impulse Responses After An Expansionary Monetary Policy Shock(Continued)



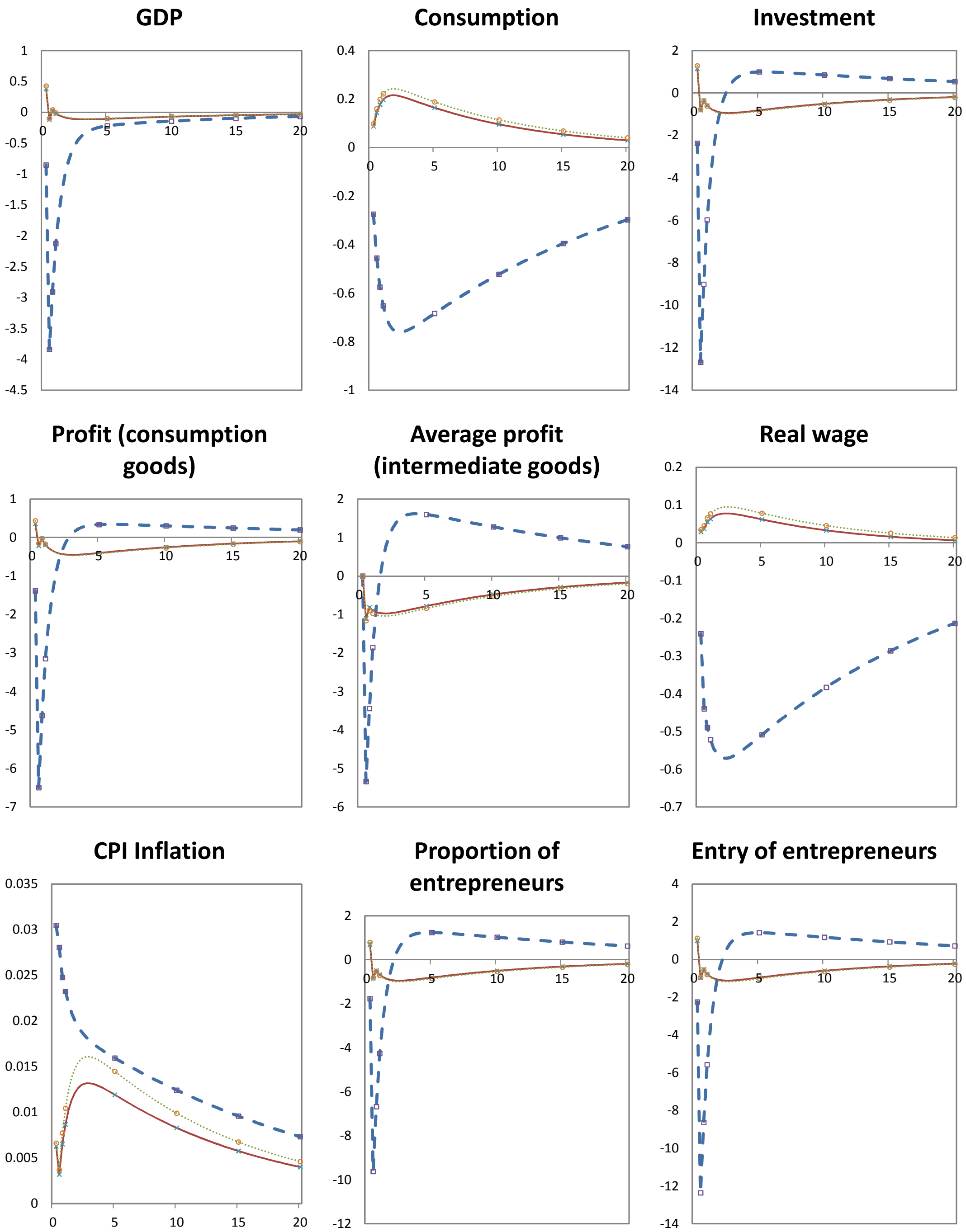
Notes: The variable on the horizontal axis is the number of years after the shock. The responses are normalized so that one denotes one percent deviation from the steady-state level. The dashed curves with square markers correspond to the responses to the shocks under the interest rate smoothing rule without reacting to output fluctuations. The solid curves with cross markers correspond to the responses under the interest smoothing rule reacting to output fluctuations. The dotted curves with round markers correspond to the responses under the forward-looking Taylor rule.

Impulse Responses After An Expansionary Monetary Policy Shock (Continued)



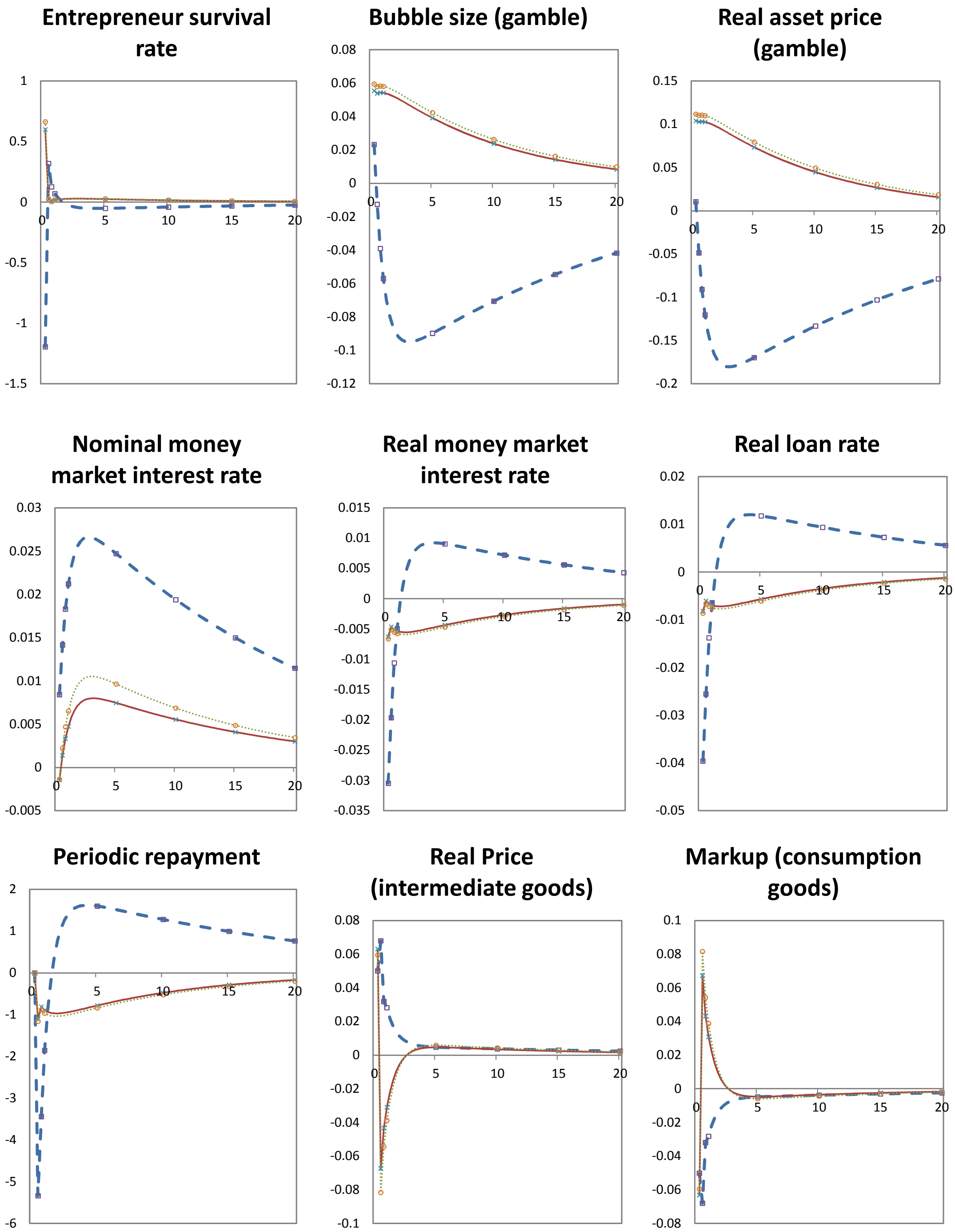
Notes: The variable on the horizontal axis is the number of years after the shock. The responses are normalized so that one denotes one percent deviation from the steady-state level. The dashed curves with square markers correspond to the responses to the shocks under the interest rate smoothing rule without reacting to output fluctuations. The solid curves with cross markers correspond to the responses under the interest smoothing rule reacting to output fluctuations. The dotted curves with round markers correspond to the responses under the forward-looking Taylor rule.

Figure A2: Impulse Responses After A Negative Aggregate Productivity Shock



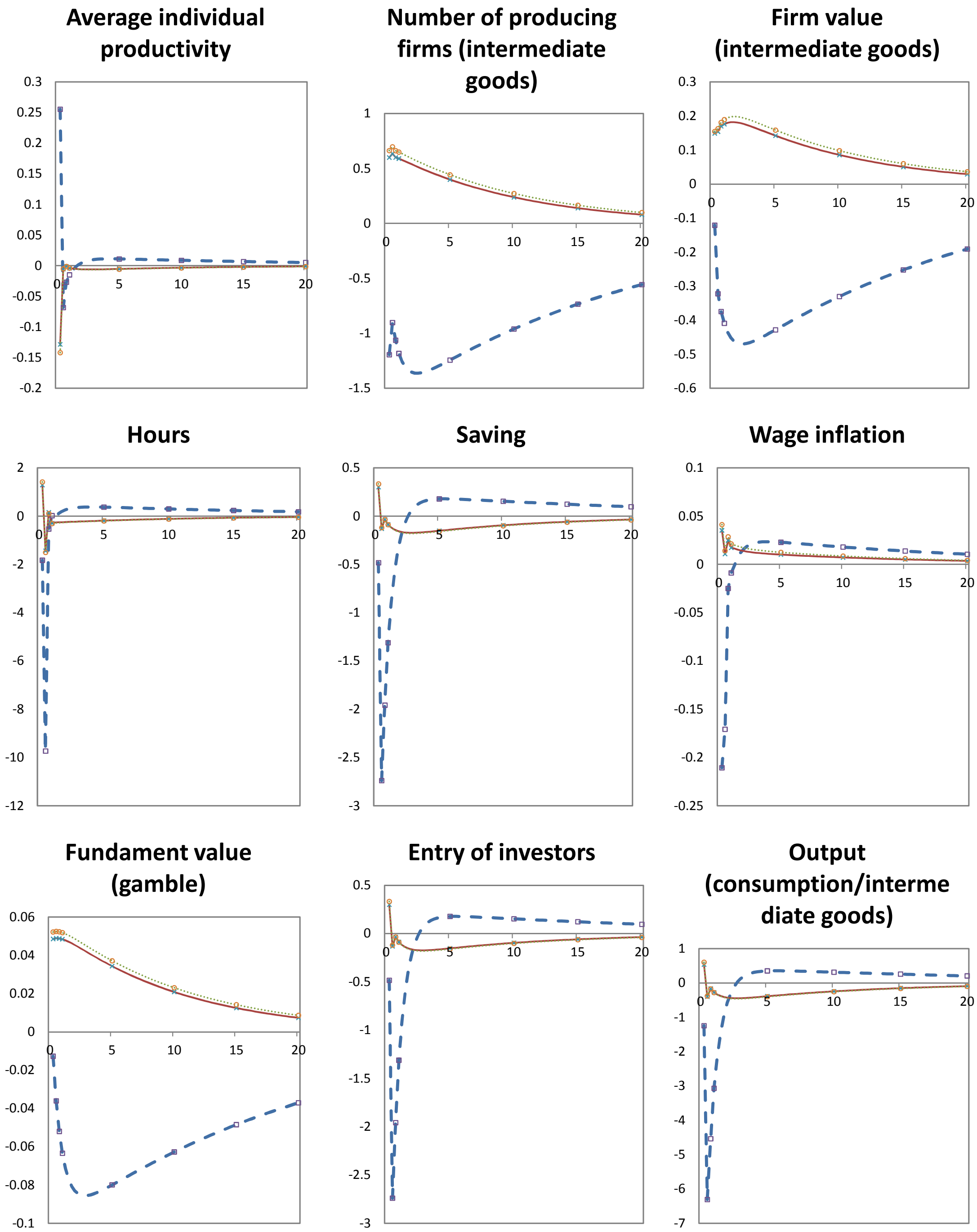
Notes: The variable on the horizontal axis is the number of years after the shock. The responses are normalized so that one denotes one percent deviation from the steady-state level. The dashed curves with square markers correspond to the responses to the shocks under the interest rate smoothing rule without reacting to output fluctuations. The solid curves with cross markers correspond to the responses under the interest smoothing rule reacting to output fluctuations. The dotted curves with round markers correspond to the responses under the forward-looking Taylor rule.

Impulse Responses After A Negative Aggregate Productivity Shock (Continued)



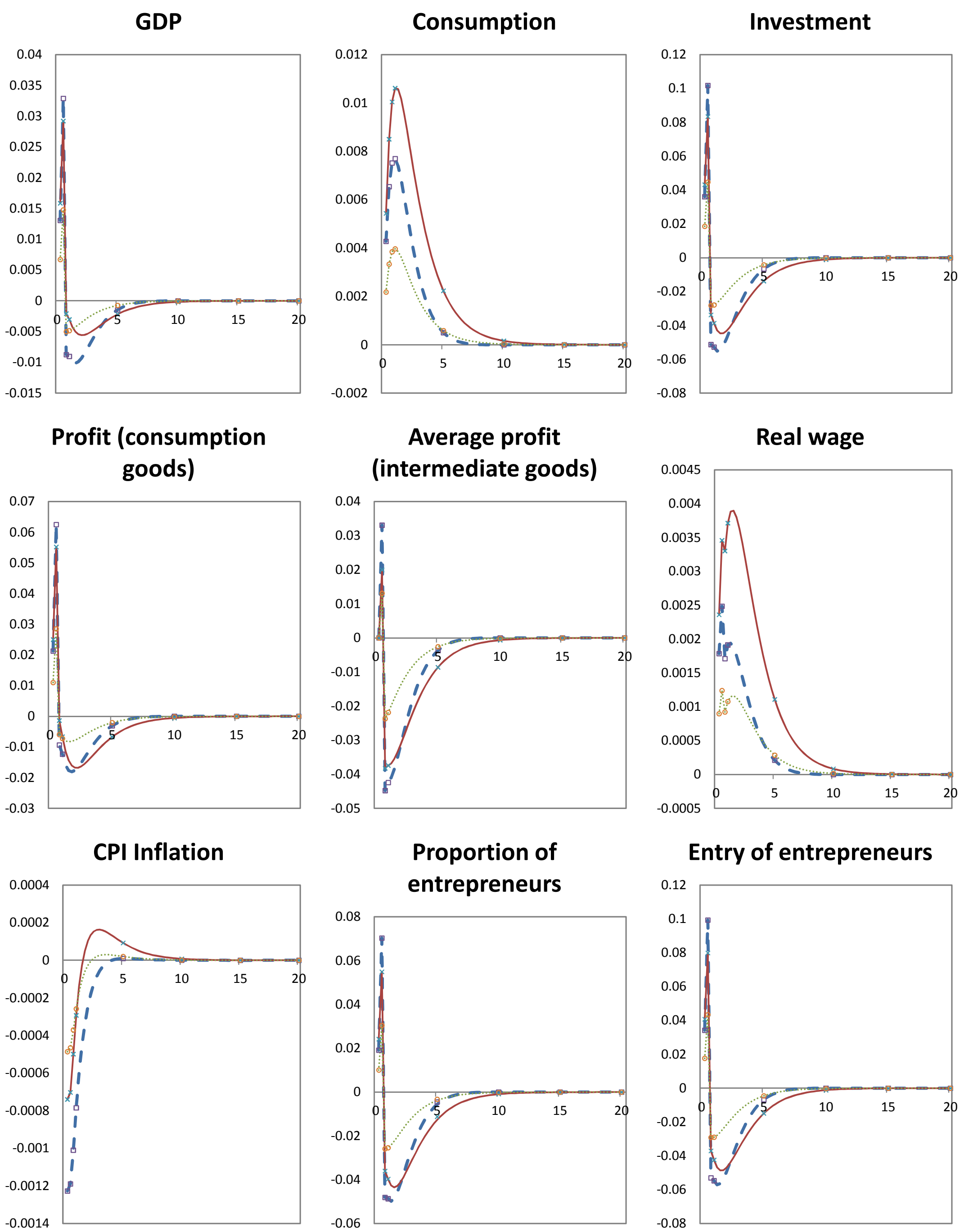
Notes: The variable on the horizontal axis is the number of years after the shock. The responses are normalized so that one denotes one percent deviation from the steady-state level. The dashed curves with square markers correspond to the responses to the shocks under the interest rate smoothing rule without reacting to output fluctuations. The solid curves with cross markers correspond to the responses under the interest smoothing rule reacting to output fluctuations. The dotted curves with round markers correspond to the responses under the forward-looking Taylor rule.

Impulse Responses After A Negative Aggregate Productivity Shock (Continued)



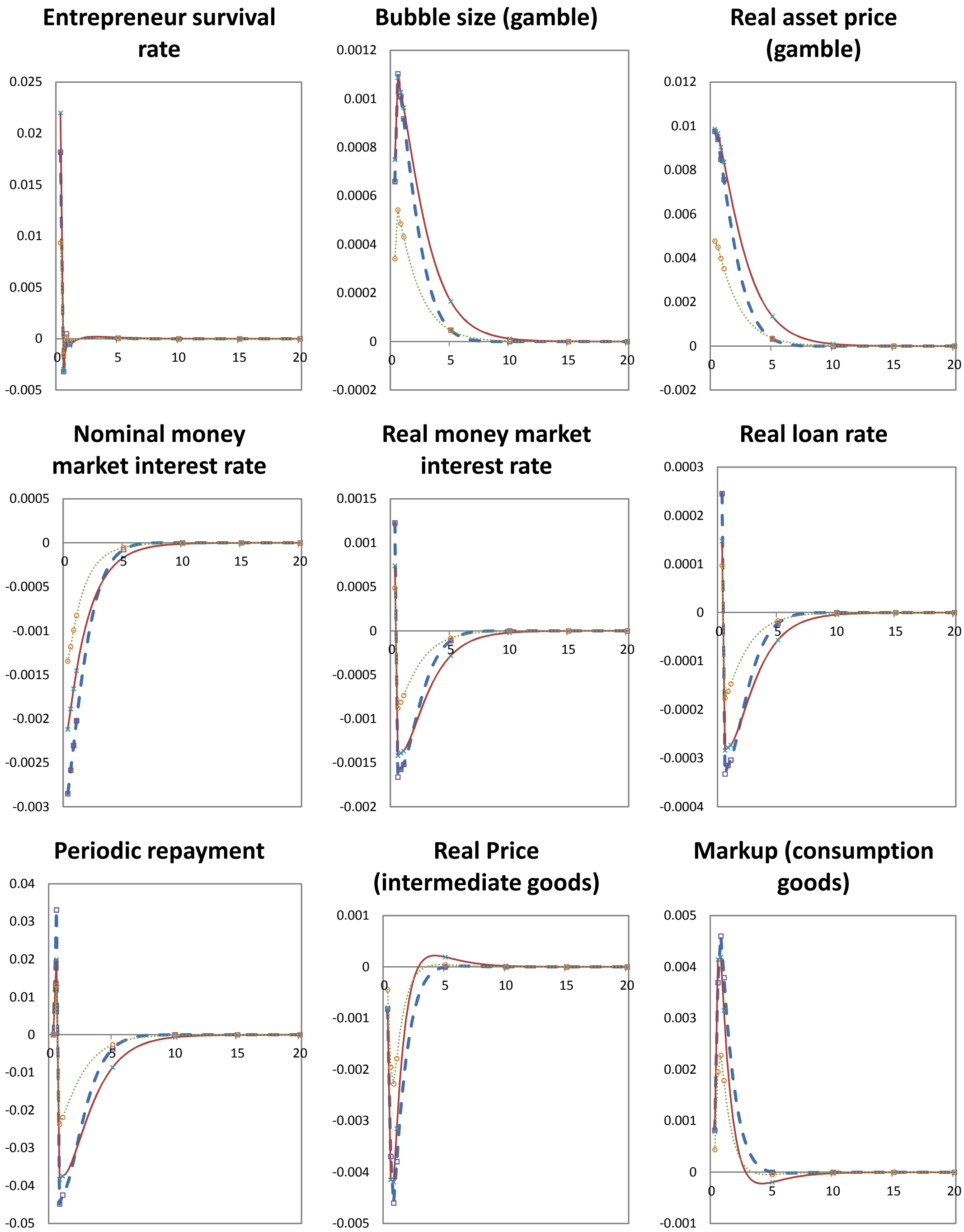
Notes: The variable on the horizontal axis is the number of years after the shock. The responses are normalized so that one denotes one percent deviation from the steady-state level. The dashed curves with square markers correspond to the responses to the shocks under the interest rate smoothing rule without reacting to output fluctuations. The solid curves with cross markers correspond to the responses under the interest smoothing rule reacting to output fluctuations. The dotted curves with round markers correspond to the responses under the forward-looking Taylor rule.

Figure A3: Sticky Interest Rate Passthrough and Impulse Responses After An Expansionary Monetary Policy Shock



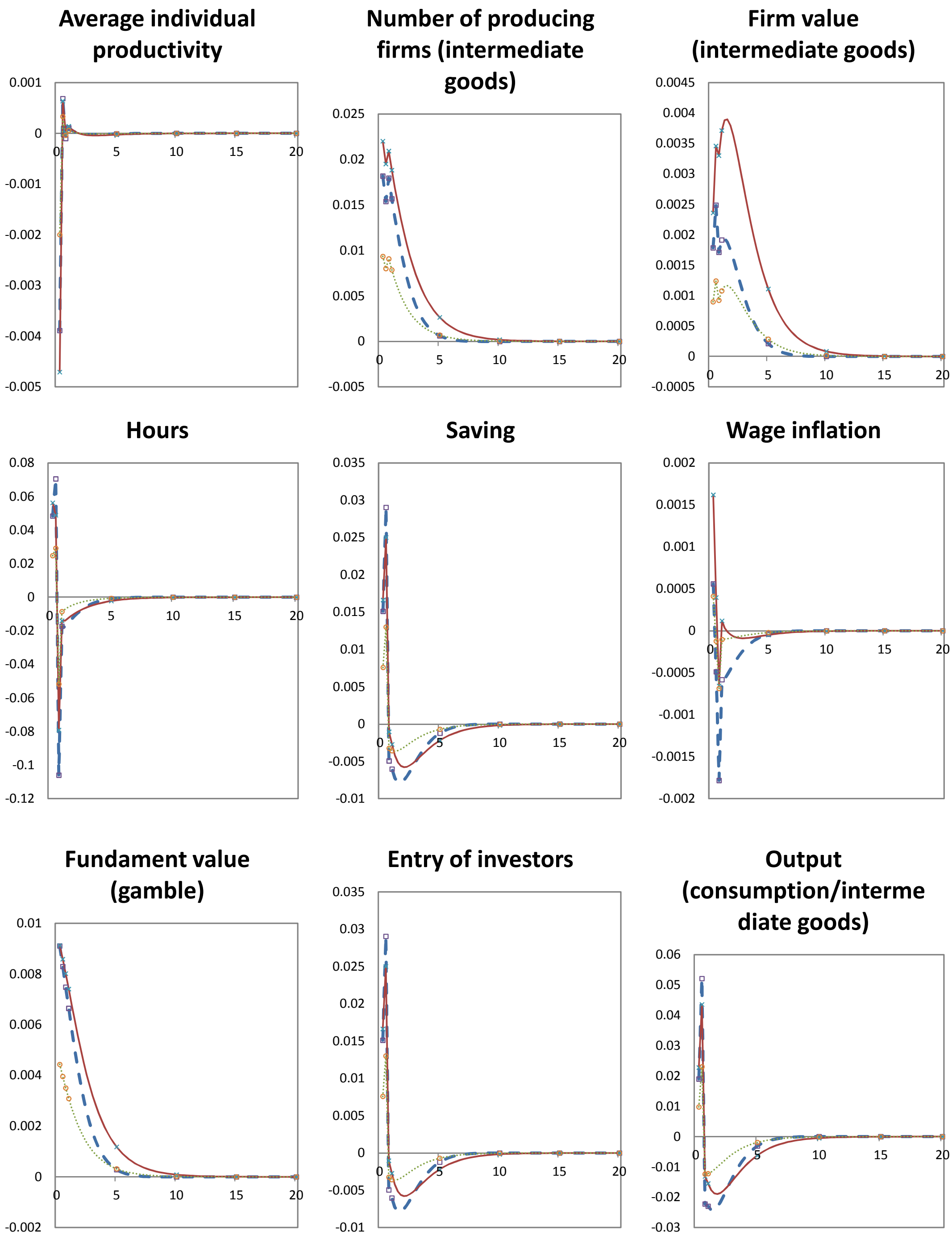
Notes: The variable on the horizontal axis is the number of years after the shock. The responses are normalized so that one denotes one percent deviation from the steady-state level. The dashed curves with square markers correspond to the responses to the shocks under the interest rate smoothing rule without reacting to output fluctuations. The solid curves with cross markers correspond to the responses under the interest smoothing rule reacting to output fluctuations. The dotted curves with round markers correspond to the responses under the forward-looking Taylor rule.

Sticky Interest Rate Passthrough and Impulse Responses After An Expansionary Monetary Policy Shock (Continued)



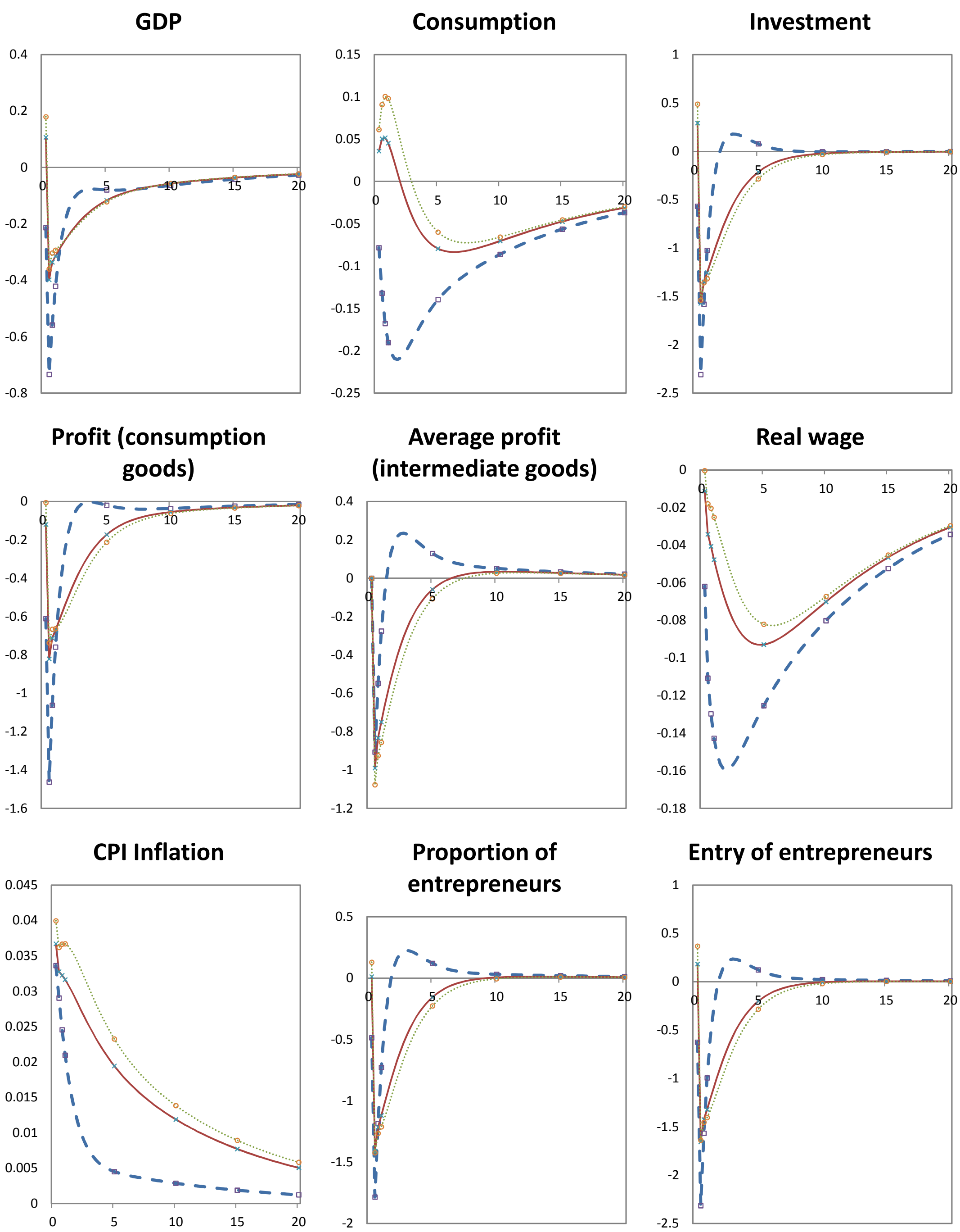
Notes: The variable on the horizontal axis is the number of years after the shock. The responses are normalized so that one denotes one percent deviation from the steady-state level. The dashed curves with square markers correspond to the responses to the shocks under the interest rate smoothing rule without reacting to output fluctuations. The solid curves with cross markers correspond to the responses under the interest smoothing rule reacting to output fluctuations. The dotted curves with round markers correspond to the responses under the forward-looking Taylor rule.

Sticky Interest Rate Passthrough and Impulse Responses After An Expansionary Monetary Policy Shock (Continued)



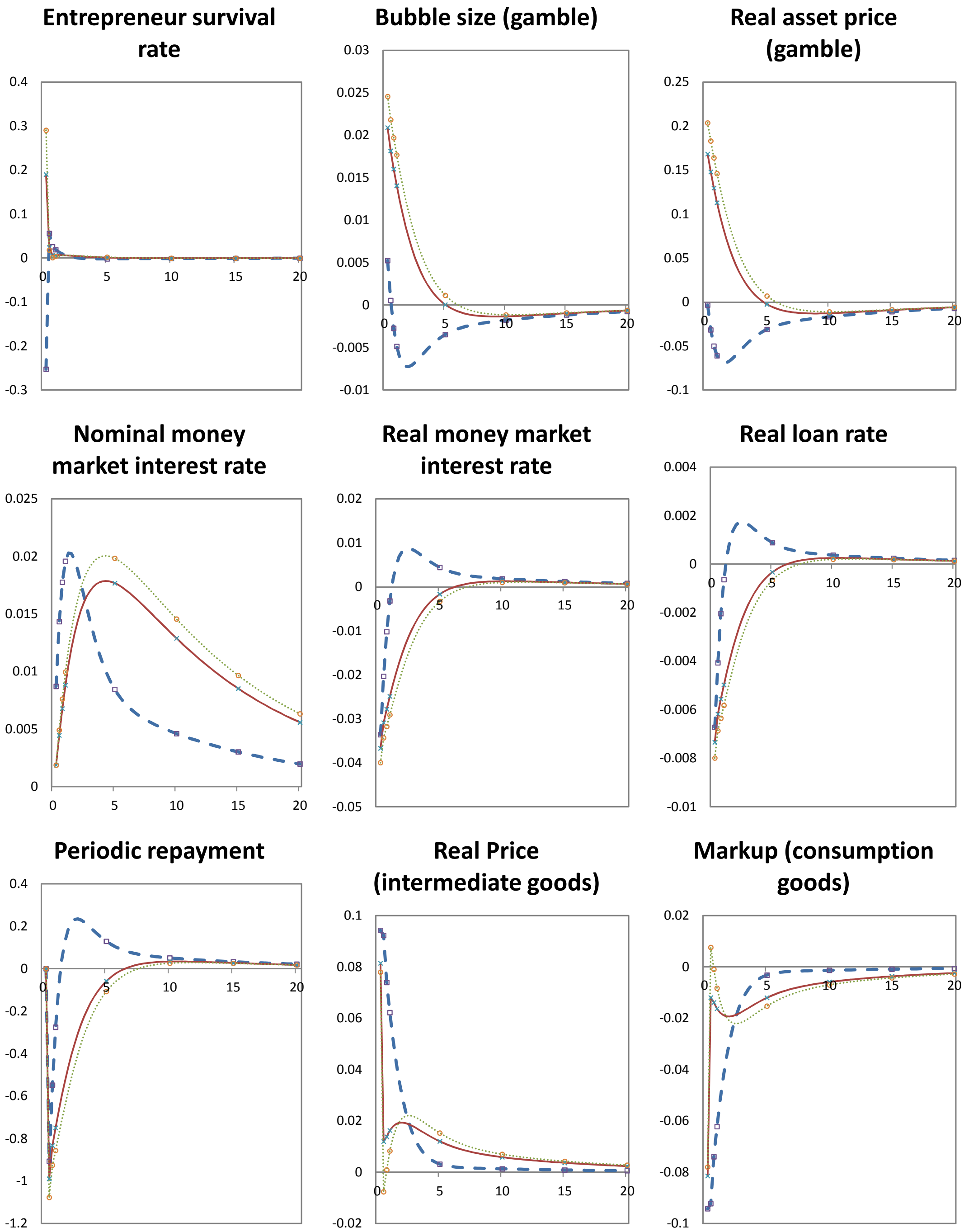
Notes: The variable on the horizontal axis is the number of years after the shock. The responses are normalized so that one denotes one percent deviation from the steady-state level. The dashed curves with square markers correspond to the responses to the shocks under the interest rate smoothing rule without reacting to output fluctuations. The solid curves with cross markers correspond to the responses under the interest smoothing rule reacting to output fluctuations. The dotted curves with round markers correspond to the responses under the forward-looking Taylor rule.

Figure A4: Sticky Interest Rate Passthrough and Impulse Responses After A Negative Aggregate Productivity Shock



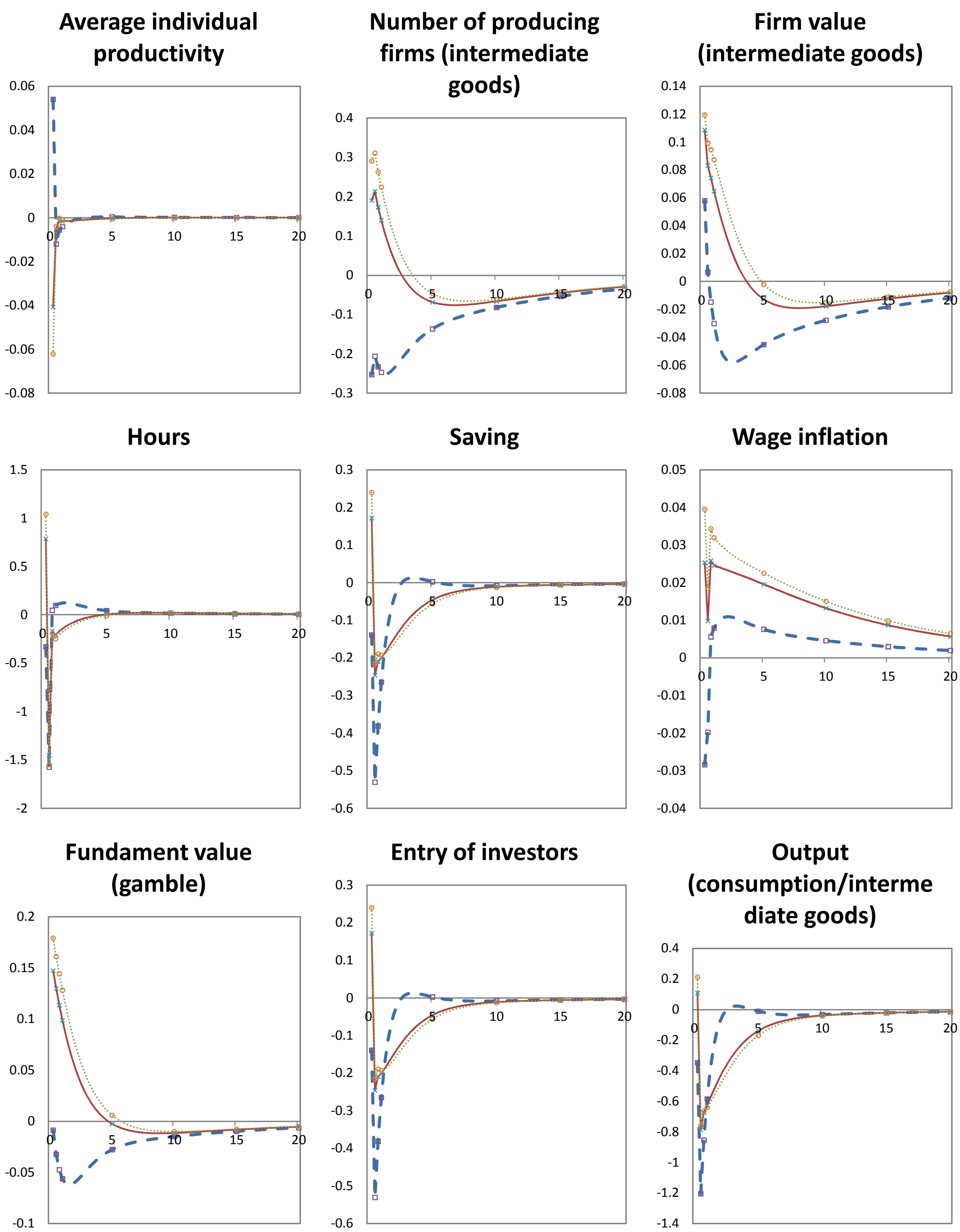
Notes: The variable on the horizontal axis is the number of years after the shock. The responses are normalized so that one denotes one percent deviation from the steady-state level. The dashed curves with square markers correspond to the responses to the shocks under the interest rate smoothing rule without reacting to output fluctuations. The solid curves with cross markers correspond to the responses under the interest smoothing rule reacting to output fluctuations. The dotted curves with round markers correspond to the responses under the forward-looking Taylor rule.

Sticky Interest Rate Passthrough and Impulse Responses After A Negative Aggregate Productivity Shock (Continued)



Notes: The variable on the horizontal axis is the number of years after the shock. The responses are normalized so that one denotes one percent deviation from the steady-state level. The dashed curves with square markers correspond to the responses to the shocks under the interest rate smoothing rule without reacting to output fluctuations. The solid curves with cross markers correspond to the responses under the interest smoothing rule reacting to output fluctuations. The dotted curves with round markers correspond to the responses under the forward-looking Taylor rule.

Sticky Interest Rate Passthrough and Impulse Responses After A Negative Aggregate Productivity Shock (Continued)



Notes: The variable on the horizontal axis is the number of years after the shock. The responses are normalized so that one denotes one percent deviation from the steady-state level. The dashed curves with square markers correspond to the responses to the shocks under the interest rate smoothing rule without reacting to output fluctuations. The solid curves with cross markers correspond to the responses under the interest smoothing rule reacting to output fluctuations. The dotted curves with round markers correspond to the responses under the forward-looking Taylor rule.

Chapter 4

TRADE OPENNESS AND THE PHILLIPS CURVE: THE NEGLECTED HETEROGENETIY AND ROBUSTNESS OF EMPIRICAL EVIDENCE[58]

4.1. Introduction

Theoretical models (Romer, 1993; Lane, 1997; Razin and Loungani, 2005; Daniels and VanHoose, 2006) suggest that increased trade openness[59] tends to reduce a country's trend inflation rate by affecting the slope of a country's Phillips curve. Are these models good explanations for the reduction in national inflation rates of OECD countries in the last few decades? To answer this question, we have to first answer the question of whether trade openness has significantly affected the slope of the Phillips curve in OECD countries.

A popular approach used in the literature (Temple, 2002; Daniels et al., 2005; Daniels and VanHoose, 2009; Badinger, 2009) to test the trade openness-Phillips curve correlation is a cross-country regression in which the parameters of the equation are assumed to be homogeneous across countries. In this chapter, we argue that the validity of the parameter homogeneity assumption is not guaranteed from a theoretical perspective and that the econometric analysis should start without an *a priori* parameter homogeneity assumption. Following such a principle, we start our econometric analysis with country-

[58] This chapter is coauthored with Sylvester Eijffinger.

[59] In this chapter, we measure trade openness by total imports and exports divided by gross domestic product (GDP).

specific time series analysis.

In a recent time series study, Ihrig et al. (2010) find no effect of trade openness on the slope of the Phillips curve in advanced economies. This result is consistent with the previous cross-country regression results. Although Daniels et al. (2005) find that trade openness significant, their more recent study (Daniels and VanHoose, 2009) find trade openness has no effect. However, using the same sample (the sample covers Australia, Canada, France, Italy, Japan, Netherlands, Sweden, United Kingdom and United States from the year 1977 to 2007) we show that the "no effect" result of Ihrig et al. (2010) is due to model misspecifications. More specifically, there are both redundant and omitted explanatory variables in their empirical model. Additionally, potential structural breaks in the inflation persistence of the sample countries are not considered. Correcting for these misspecifications, we find that trade openness has significant effects on the slope of the Phillips curve in several major industrial countries.

Badinger (2009) suggests that omitting the interaction between the degree of financial openness[60] and the output gap in the regression can cause an endogeneity problem. More specifically, trade and financial openness are highly correlated. If both have significant effects on the slope of the Phillips curve, omitting one of them will cause an omitted variable bias. He finds both financial openness and trade openness are significant in the sample of developing countries while these variables are not significant in the sample of industrialized countries. By contrast, we find trade and/or financial openness significant in several major industrial countries. The important reason is that we relax the parameter homogeneity assumption of Badinger (2009). Actually, among the sample countries (Canada, France, Italy, Sweden and the United States) where we find openness significant, both the size and sign of the effects differ.

Some authors (Ihrig et al., 2010; Ball, 2006) use panel data methods to test the trade openness-Phillips curve correlation and find that trade openness has no significant effect on the slope of the Phillips curve in industrialized countries. The "no effect" result is again associated with the parameter homogeneity assumption. More specifically, they assume that the coefficients of the explanatory variables are the same across countries. We find that this assumption is not valid. The seemingly unrelated regressions (SUR) estimator allows both the intercepts and the slopes of the country-specific models to be

[60]The degree of financial openness is measured by total foreign assets and liabilities divided by GDP.

heterogeneous across countries. Therefore, it is free of the bias from the false parameter homogeneity assumption. We estimate the panel by SUR and find that trade openness has significantly changed the slope of the Phillips curve in several important industrialized countries. IMF (2006) also uses SUR to estimate the effect of trade openness on the slope of the Phillips curve in several industrial countries. Similar to our finding, the IMF study shows that trade openness has played a significant role. However, it finds that trade openness uniformly flattens the Phillips curve in the sample of countries while we find the effects of trade openness differ in sign across countries. Although the IMF (2006) study does allow trade openness to have heterogeneous effects across countries, it restricts the heterogeneity by assuming that trade openness affects the slope of each country's Phillips curve through a common multiplicative term. This restriction is not guaranteed to be valid from a theoretical perspective (see Section 4.2). Our results suggest that the effects of trade openness can be significantly different in both sign and size. Hence, restricting the cross-country heterogeneity when estimating the empirical model may not be a good idea.

The chapter proceeds as follows. Section 2 provides the theoretical foundation of why the parameter homogeneity assumption may be problematic. Section 3 performs country-specific time series analysis for our sample OECD countries. Section 4 extends the analysis to a panel data setting. Section 5 concludes.

4.2. Theoretical background

Due to differences in modeling strategies and behavioral assumptions, previous theoretical models on the trade openness-Phillips curve correlation give different predictions on the effect of trade openness on the slope of the Phillips curve. The models of Romer (1993) and Lane (1997) predict that an increase in trade openness steepens the Phillips curve, while the models of Razin and Loungani (2005) and Daniels and VanHoose (2006) predict that an increase in trade openness flattens the Phillips curve. As a consequence, previous cross-country empirical studies (Temple, 2002; Daniels et al., 2005; Daniels and VanHoose, 2009; Badinger, 2009) use the sign and statistical significance of estimated trade openness-Phillips curve correlation to test the empirical relevance of various theoretical models. For example, Daniels et al. (2005) find that an increase in trade openness flattens the Phillips curve in the OECD, and they take this finding as a support for the theoretical model of Daniels and VanHoose (2006).

A common restriction imposed in the previous cross-country empirical studies is parameter homogeneity across countries. However, the validity of the parameter homogeneity restriction is not guaranteed from a theoretical perspective. An increase in trade openness can flatten the Phillips curve in some countries while steepen the Phillips curve in other countries. In a New Keynesian model with Kimball (1995) preferences, Sbordone (2007) shows that the open economy Phillips curve takes the following form:

$$\pi_t = \beta E_t(\pi_{t+1}) + \lambda \widehat{rmc}_t, \tag{4.1}$$

where π_t is the inflation rate, β is the time discount factor, rmc_t is the economy-wide average real marginal cost of production and E_t is expectations operator. Throughout this chapter, we shall use a hat over a variable to denote the deviation of this variable from its steady-state level, so $\widehat{rmc}_t$ is the deviation of real marginal cost from its steady state level. λ is the slope of the Phillips curve which is determined by a number of factors. More specifically,[61]

$$\lambda = \kappa[1 + \bar{\theta}(\bar{\epsilon}_\mu + \bar{s}_y)]^{-1}, \tag{4.2}$$

where $\kappa \in (0, 1)$ is a function of the time discount factor (β) and the degree of price stickiness, $\bar{\theta}$ is the steady-state price elasticity of demand, $\bar{\epsilon}_\mu$ is the steady-state elasticity of the firm's desired markup to its market share, $\bar{s}_y$ is the steady-state elasticity of the firm's marginal cost to its own output.[62] Intuitively, an increase in the firm's own price will raise its relative price due to the fact that some of the firms cannot change their prices in the current period. A higher demand elasticity means a higher loss of the firm's market share due to an increase in its relative price. Similarly, if $\bar{\epsilon}_\mu$ is higher, a decrease in market share will reduce the firm's desired markup to a larger extent, so the firm will be less willing to increase its relative price. Therefore, an increase in $\bar{\theta}$ or $\bar{\epsilon}_\mu$ will weaken the firm's incentive to increase its own price after an increase in the average marginal cost. Moreover, if there are firm-specific inputs, the firm's own marginal cost not only

[61] See Sbordone (2007) for a derivation.

[62] Throughout this chapter, we shall use a bar over a variable to denote its steady state.

depends on the economy-wide average marginal cost, but also depends on its own output. If the firm's marginal cost is increasing in its own output, its desired price increase after an increase in the economy-wide average marginal cost will be limited. This relationship is because the firm will benefit from the decline in the firm-specific marginal cost after its relative price increases and market share declines. Therefore, the firm's price increases less after an increase in the economy-wide average marginal cost if the elasticity of the firm's marginal cost to its own output in the steady state ($\bar{s}_y$) is higher.

Sbordone (2007) argues that an increase in the degree of trade openness can be taken as an increase in the number of competing firms in the market. An increase in the number of competing firms will reduce each firm's market share, which increases the price elasticity of demand $\bar{\theta}$. This result is because with Kimball (1995) preferences, the demand curve faced by the individual firm is downward sloping so that $\bar{\theta}$ is a decreasing function of the market share.[63] Therefore, an increase in the degree of trade openness increases $\bar{\theta}$.

Sbordone (2007) assumes that the logarithm of the price markup is a convex function of the logarithm of the market share.[64] Under this assumption, it is easy to show that $\bar{\epsilon}_\mu$ is decreasing in the number of competing firms. To see this, let us denote the price markup by μ, the market share by x. ϵ_μ is defined as $\frac{\partial log\mu(x)}{\partial logx}$. Its derivative with respect to $logx$ is $\frac{\partial(\partial log\mu(x)/\partial logx)}{\partial logx}$, which is positive due to the convexity assumption. Therefore, ϵ_μ is an increasing function of the logarithm of the market share. Since the logarithmic function is an increasing function, ϵ_μ is an increasing function of the market share as well. An increase in the number of competing firms decreases each firm's market share, and

[63]See Kimball (1995).

[64]The motivation for this assumption is that the logarithm of the price markup cannot be a concave function of the logarithm of the market share. Because the price elasticity of demand is a decreasing function of the market share and the price markup is a decreasing function of the price elasticity of demand, the price markup is an increasing function of the market share. Since the logarithmic transformation is a monotonically increasing transformation, the logarithm of the price markup is also an increasing function of the logarithm of the market share. This means that as the logarithm of the market share declines, the logarithm of the price markup also declines. As the market share approaches zero, the logarithm of the market share approaches negative infinity. Due to concavity, the decline in the logarithm of the price markup will be faster than the decline in the logarithm of the market share. Therefore, the logarithm of the price markup approaches negative infinity as the logarithm of the market share approaches negative infinity. This cannot happen because the price markup is larger than one, which means that its logarithm is always positive. Note that this reasoning only proves that the logarithm of the price markup is a convex function of the logarithm of the market share for small values of the market share. However, in a monopolistically competitive market with symmetric firms, it is reasonable to believe that the market share for each firm is very small. Thus, it is convenient to assume global convexity.

therefore reduces ϵ_μ. This property also holds in the steady state, so $\bar{\epsilon}_\mu$ is a decreasing function of the number of competing firms. Hence, an increase in the degree of trade openness decreases $\bar{\epsilon}_\mu$.

Sbordone (2007) shows that the sign of the derivative of $\bar{s}_y$ with respect to the degree of trade openness is ambiguous, depending on a number of factors (the Frisch elasticity of labor supply, the labor share, and the fixed cost of production). Therefore, the net effect of a change in the degree of trade openness on the slope of the Phillips curve is ambiguous, depending on the relative changes in $\bar{\theta}$, $\bar{\epsilon}_\mu$, $\bar{s}_y$ after a change in trade openness. The net effects of trade openness on the slope of the Phillips curve will differ in size and/or sign across countries if those relative changes after a change in trade openness differ across countries, which implies that a parameter homogeneity restriction in the econometric analysis is potentially problematic. For this reason, we start our econometric analysis in Section 3 with individual time series analysis which does not impose the parameter homogeneity restriction.

4.3. Country-specific time series analysis

4.3.1. The empirical model of Ihrig et al. (2010)

We start our econometric analysis with the following backward-looking Phillips curve model of Ihrig et al. (2010) for $t = 1, ..., T$:

$$\pi_t^c = \phi_0 + \phi_1 \pi_{t-1}^c + \phi_2 \hat{y}_t + \phi_3 \alpha_t \hat{y}_t + \phi_4 p_t^e + \phi_5 p_t^f + \phi_6 p_t^m + \phi_7 p_t^m * Mshare_t + \varepsilon_t, \quad (4.3)$$

where π_t^c is the core consumer price index (CPI) inflation rate which is used by Ihrig et al. (2010) as a measure of the inflation rate (π_t) in the theoretical model; α_t is the trade openness measured as total imports and exports divided by GDP; $\hat{y}_t$ is the output gap; p_t^e, p_t^f, p_t^m are the deviation of energy, food and import price changes from the last period core CPI inflation rate, respectively; $Mshare_t$ is import as a share of GDP; and ε_t is a normal i.i.d. error term.[65] As surveyed by Gordon (2011), there is a debate on the

[65] Previous empirical studies (Culver and Papell, 1997; Basher and Westerlund, 2007) find that the inflation rate is stationary, so our econometric analysis is performed under the assumption that the inflation rate is stationary.

empirical modeling of inflation expectations. Some economists use the backward-looking assumption as in equation (4.3) while others use a forward-looking assumption. We adopt the backward-looking assumption because the estimation of the forward-looking model involves instrumental variables and the results are subject to weak instrument problems (Kleibergen and Mavroeidis, 2009; Nason and Smith, 2008). The focus of this chapter is on the validity of the parameter homogeneity assumption in the previous empirical models. Hence, it is better to separate the focus issue from the instrument quality issue. Moreover, previous studies (Ihrig et al., 2010; Ball, 2006; IMF, 2006) on the openness-Phillips curve correlation typically adopt the backward-looking assumption. Therefore, it is easier to compare the results if we use the same assumption. Another debate related to the specification of equation (4.3) is on whether or not one should include $p_t^e, p_t^f, p_t^m, Mshare_t$ into the empirical model. Those variables are called the cost push terms. Ball (2006) argues that those terms should not be included in the Phillips curve estimation. This argument is rooted in the theoretical model of Ball and Mankiw (1995) in which smooth relative price changes, such as smooth changes in the price of energy, food and import goods relative to the general price level, do not affect the general price level. The empirical validity of that model, however, is challenged by Bryan and Cecchetti (1999). Gordon (2011) justifies the role of relative price changes by price rigidity in sectors which are not subject to the relative price shocks. Monacelli (2005) suggests that in an open economy with incomplete exchange rate pass-through, additional cost push terms must be added to the Phillips curve if the output gap is used to measure $\widehat{rmc}_t$. Batini et al. (2005) suggest that the signs of the cost push terms in the Phillips curve can be either positive or negative, depending on the elasticity of material inputs with respect to gross output. Due to the theoretical ambiguity, we do not impose any sign or size restriction on the cost push terms and will apply the general-to-specific model selection strategy to eliminate redundant variables in section 4.3.2. For now, we just take equation (4.3) as given and try to replicate the results of Ihrig et al. (2010).

We use similar data source as Ihrig et al. (2010). However, the frequency of our model is annual rather than quarterly because one of the important control variables, that is, financial openness, in Section 4.3.3, is sampled at the annual frequency (see Table 4.1 and 4.2 for more information about the data). The sample period is 1977-2007, which covers the sample period of Ihrig et al. (2010), 1977-2005. The Ihrig et al.

(2010) sample consists of eleven OECD countries. Our sample takes nine out of these eleven countries.[66] While Ihrig et al. (2010) include more lags of core inflation and the cost push terms in their quarterly model, we include only the first lag of the core inflation rate in equation (4.3) to save degrees of freedom. Serial independence tests (Table 4.3) suggest that including the first lag of the core inflation rate is enough to eliminate serial correlation in the model for most sample countries. As we shall see later, in more robust model specifications, doing so is enough to eliminate serial correlation in the model for all sample countries. We lag the independent variables by one period to avoid potential bias from reverse causality. An additional reason to lag the cost push terms is to avoid collinearity between them and the lagged dependent variable.

Table 4.3 presents the estimation results of the Ihrig et al. (2010) model with our annual dataset. Similar to Ihrig et al. (2010), we find no significant effect of trade openness in most sample countries. However, it is also noticeable that not only the interaction term between trade openness and the output gap, but also many other explanatory variables are not statistically significant. In addition, specification tests suggest that the serial independence and normality assumption on the error terms are violated in several sample countries, which can bias the inference. If serial correlation and non-normality are the only problems, one can easily correct for them by adding more lags of inflation to the model and removing outliers which cause the non-normality. Ihrig et al. (2010) actually carry out those corrections and still get no significant role of trade openness in their model. However, there are more problems with equation (4.3) than just serial correlation and non-normality.

4.3.2. Toward a more robust specification

In this subsection, we consider the following potential problems associated with equation (4.3). First, as argued by Benati (2008), inflation persistence may have changed over time due to institutional changes. We use the Andrews (1993) test to check for potential structural breaks in the inflation persistence of the sample countries. Even if we have prior expectations that some events may change the process of inflation, how long does it take for the effects of the events to be fully absorbed is uncertain. The advantage of

[66] Ihrig et al. (2010) measure the output gap of the sample countries by the OECD output gap estimates. When the OECD output gap estimates are missing, as is the case for Switzerland, they use the Hodrick and Prescott (HP) filtered output gap instead. We only include countries with the OECD output gap estimates in our sample for consistency reasons.

the Andrews (1993) test over the standard Chow test is that it does not require us to specify an arbitrary year for the structural break. Therefore, it avoids the bias from a misspecified break point.[67]

In six out of nine countries, i.e., Australia, Canada, France, The Netherlands, Sweden, United States, we find a significant structural break in inflation persistence (see Table 4.3). This suggests that we need to model the potential structural break in the inflation persistence to get the right inference. Second, there may be omitted variables in equation (4.3). We control for four potentially omitted variables in our augmented version of the Ihrig et al. (2010) model.[68] The first two variables are interaction terms between the logarithm of GDP (log GDP), the logarithm of population (log population) and the output gap. Log GDP and log population are proxies for the country size. We include the interaction term between country size and the output gap into the model because Sbordone (2007) suggests that market size, that is, the number of competing firms, can affect the slope of the Phillips curve. As we discussed in Section 4.2, the expected sign of a proxy for the country size effect is ambiguous. The third additional control variable we add to the Ihrig et al. (2010) model is an interaction term between trend inflation (proxied by the HP filtered trend of core inflation rate) and the output gap. The reason for adding this control variable is that the state dependent pricing models (Ball et al., 1988; Bakhshi et al., 2007) suggest that the slope of the Phillips curve changes with the degree of price stickiness which is affected by the trend inflation rate. Last but not least, we add the "global inflation" variable defined by Ciccarelli and Mojon (2010). These authors find that there is a common factor in the OECD countries' national inflation rates and they call this common factor "global inflation". Ciccarelli and Mojon (2010) suggest that a simple cross-country average of 22 OECD countries.[69] fits the "global inflation" well, so we follow them and proxy global inflation by the simple cross-country average of the 22 OECD countries considered by Ciccarelli and Mojon (2010).

[67]Compared to a Chow test with correctly specified structural breaks, the Andrews (1993) test is less efficient. To test the robustness of results, we perform the Chow test with break points detected by the Andrews (1993) test. The results are consistent with the Andrews (1993) test results.

[68]In previous cross-country studies, researchers also control for the effect of central bank independence (CBI) on the slope of the Phillips curve. Although CBI has changed in our sample countries, yearly variation is limited. We tried to control for the effect of CBI by adding an interaction term between the output gap and the dynamic GMT index constructed by Arnone and Romelli (2012). It turns out that this additional interaction term is not significant in the regressions.

[69]The 22 OECD countries are Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Japan, Luxembourg, New Zealand, Norway, Portugal, Sweden, Switzerland, Spain, United Kingdom, United States, and the Netherlands.

Our modeling strategy is as follows: we always include a constant term, lagged core inflation rate, the output gap and the interaction term between trade openness and the output gap in the model as our focus variables. Other variables are taken as control variables in the model. We remove them from the model if they are detected to be redundant by the F-test. We introduce an interaction term between a break dummy and the lagged dependent variable into the model whenever a structure break in the inflation persistence is identified by the Andrews (1993) test. Moveover, we eliminate outliers from the model by adding dummy variables which take the value one in the outlier year and zero otherwise.[70]

Table 4.4 summarizes the estimation results of our models. First note that equation (4.3) clearly omits some important control variables. Particularly, global inflation is significant in seven out of the nine sample countries. This is consistent with the finding by Ciccarelli and Mojon (2010) that global inflation accounts for a large proportion of the variance in the inflation rates of the OECD countries. The degree of inflation persistence significantly declines in recent years in three of the sample countries (Australia, Canada, and the United States).[71] The estimated coefficients of our variable of interest, the interaction term between trade openness and the output gap, differ in sign across countries. More importantly, they are statistically significant in two of the major industrialized countries, that is, Canada and France. More specifically, the estimated coefficient of the interaction term between trade openness and the output gap is negative in Canada, suggesting that trade openness flattens the Phillips curve. By contrast, there is a steepening effect of trade openness in France.[72]

4.3.3. Controlling for financial openness

So far, we have only considered the effect of trade openness on the slope of the Phillips curve. Badinger (2009) suggests that financial openness can also affect the slope of

[70] See the table notes for the detected outlier years.

[71] The break points are 1991, 1987 and 1982, respectively. The break points in other countries detected in Table 1 are no longer significant in the current model augmented by additional control variables.

[72] Brambor et al. (2006) argue that in an econometric model with an interaction term, say XZ, the magnitude and statistical significance of the interaction term are not very useful to understand the marginal effect of X on the dependent variable Y. In this chapter, the magnitude and statistical significance of the interaction term do matter. The slope of the Phillips curve is the marginal effect of the output gap on the inflation rate. However, the slope itself is not our focus. The focus is the effect of openness on the slope of the Phillips curve, which is exactly the coefficient of the interaction term between openness and the output gap.

the Phillips curve.[73] Moreover, he finds that omitting financial openness can cause an endogeneity bias in the estimation. For this reason, we add an interaction term between financial openness and the output gap into the model. As noted by Badinger (2009), trade and financial openness are highly correlated. This can cause a collinearity problem in the estimation. Following Badinger (2009), we estimate the model with the restriction that the coefficient of trade openness*output gap is the same as financial openness*output gap. Formal statistical tests (Table 4.5) support the restriction in seven out of the nine sample countries. The restriction is rejected in France and Sweden. For those two countries, we report the estimation results without the restriction (in Table 4.5). Overall, the results in Table 4.5 suggest that trade openness has significantly affected the slope of the Phillips curve in Canada, France, Italy, Sweden and the US. Moreover, the effects differ in sign. There is a flattening effect of openness in Canada, Sweden and the US[74] while there is a steepening effect in France and Italy. This result questions the validity of the parameter homogeneity assumption in a cross-country regression. Therefore, when financial openness is included, our time series regression results provide stronger evidence that trade openness has significantly affected the slope of the Phillips curve in major industrialized economies. Note that for the United States, we find that both log GDP*output gap and log Popluation*output gap are statistically significant. The estimates for their coefficients are similar in magnitude but different in sign. It is possible that the joint significance of those two variables captures the effect of income per capita. For this reason, we test the restriction that the coefficients of those two variables are the same in magnitude. It turns out that this restriction cannot be rejected (the p value of the test is 0.69). Therefore, we re-estimate the empirical model with this additional restriction for the United States and report the results in Table 4.5.

[73]Following Badinger (2009), we define the degree of financial openness as total foreign assets and liabilities divided by GDP. The data used to construct the financial openness measure are sourced from the updated and extended version of the External Wealth of Nations Mark II database developed by Lane and Milesi-Ferretti (2007).

[74]Note that although the estimated coefficient of the output gap is negative for the US, it does not suggest that there is a negative correlation between the output gap and the core inflation rate in the US. The overall effect of the output gap on the inflation rate is -130.45-0.89*(trade openness+financial openness)+7.55*(log GDP-log population)+0.53*trend inflation. The average effect over the period 1977-1991 is 0.64 while the average effect over the period 1992-2007 is 0.14, which means that the Phillips curve has become flatter in the US.

4.3.4. Robustness to the HP-filter output gap measure

We have estimated the empirical models with the OECD estimates of the output gaps. As a robustness check, we also estimate the models with the HP-filtered output gap. Table 4.6 reports the results from the models selected by the general-to-specific approach. The qualitative results almost do not change. Trade openness is found to be significant in Canada, France, Italy and the United States. More specifically, trade openness flattens the Phillips curve in Canada and the US while it steepens the Phillips curve in France and Italy. The openness measures are no longer significant in Sweden, however.

4.4. Panel data analysis

It is well-known that panel data analysis may be potentially efficiency-improving since it imposes a structure, which is extra information, on the regression. However, one has to be aware that if a false structure is imposed the estimates will be biased. Ihrig et al. (2010) and Ball (2006) estimate the effect of trade openness on the slope of the Phillips curve in panel data models. Their panel data models are estimated with the assumption that the coefficients of the explanatory variables are the same across countries. This assumption is actually not valid for the industrialized countries sample we focus on here. To see this, consider the following panel data model:

$$\pi^c_{i,t} = \delta_{0i} + \delta_1 \pi^c_{i,t-1} + \delta_2 \hat{y}_{i,t} + \delta_3 \alpha_{i,t} \hat{y}_{i,t} + \tau_1' X_{i,t} + \tau_2' W_{i,t} + \varepsilon_{it}, \tag{4.4}$$

where i is the index for country $i = 1, ..., N$, $t = 1, ..., T$, τ_1 and τ_2 are vectors of parameters the vector $X_{i,t}$ contains the cost push terms, $W_{i,t}$ contains the control variables: log population*output gap, log GDP*output gap, trend inflation*output gap and global inflation. This model nests the panel data models of Ihrig et al. (2010) and Ball (2006) as special cases. More specifically, Ihrig et al. (2010) estimate a model with $\tau_2 = 0$, and Ball (2006) estimates a model with $\tau_1 = \tau_2 = 0$. we perform the Roy-Zellner poolability test (Baltagi, 2005) for equation (4.4). The test statistics is 188.47, the p value is 0.00. Therefore, the null assumption of poolability is clearly rejected. We further test the assumption that the coefficients of the interaction term between openness and the output gap are homogeneous across countries, allowing other parameters to be heterogeneous. Again the assumption is rejected at the 5 percent level (the test statistics is 17.89 and the p value is 0.02). In sum, the parameter homogeneity assumption on the parameter of interest is rejected by our panel data model, which suggests that results from previous empirical studies with the parameter homogeneity assumption are not robust. A typical fixed effects model controls for the heterogeneity in the intercepts, but still omits the heterogeneity in the slopes. Therefore, it will generate biased results when trade openness has heterogeneous effects across countries. Note that controlling for the potential omitted variable bias is important for the poolability test. If we omit the vector $W_{i,t}$ from equation (4.4), the null of parameter homogeneity cannot be rejected (the p value is 0.68). However, variables in $W_{i,t}$ are jointly significant at the 1 percent level (the Wald test statistics is 155.06 and the p value is 0.00). This result suggests the importance of controlling for omitted variable bias. Actually, the results in Table 4.3 suggest that trade openness is not significant in most sample countries if the variables in $W_{i,t}$ are omitted. This leads to the spurious parameter homogeneity assumption found by the Roy-Zeller test when we set $\tau_2 = 0$.

Equation (4.4) does not make use of the model selection results in our country-specific models. It also does not consider the effect of financial openness. To show the robustness of our country-specific time series regression results, we pool the country-specific time series models summarized in Table 4.5 into a panel and estimate it by SUR.[75] The SUR estimator allows both the intercepts and the slopes of the country-specific models to be heterogeneous across countries. Therefore, it is free of the bias from the false parameter homogeneity assumption. The SUR estimation results are summarized in Table 4.7. The qualitative results of the country-specific time series regressions are largely unchanged. Trade openness remains to be statistically significant in four advanced industrial countries: Canada, France, Sweden and the United States. Moreover, the sign of the coefficients differ in those four countries. It is negative in Canada, Sweden and the US, suggesting a flattening effect of trade openness on the Phillips curve. It is positive in France, suggesting a steepening effect. IMF (2006) also estimates a SUR model for a group of advanced industrial countries. In the IMF study, a commonality in the effects

[75]The SUR model is estimated with the OECD output gap measures. However, the main result does not change if one uses the HP-filtered output instead.

of trade openness is assumed. More specifically, trade openness is assumed to affect the slopes of different countries' Phillips curves through a common multiplicative term. IMF (2006) finds a common flattening effect in the advanced industrial countries. From the theoretical perspective of Section 4.2, there is no reason to impose the commonality restriction. Actually, if trade openness has significantly flattened the Phillips curve in a group of sample countries while steepened the Phillips curve in the other sample countries, reporting the average effect can be rather misleading. The average effect can be zero even if trade openness has played a significant role in all sample countries because the significant effects with different signs can be averaged out.

4.5. Conclusion

In this chapter, we argue that the typical assumption of parameter homogeneity used in the empirical studies of the trade openness-Phillips curve correlation is not guaranteed to be valid from an *ex ante* theoretical perspective. We test this assumption with both time series and panel data analysis. Our results suggest that the validity of the parameter homogeneity assumption is highly questionable. When the parameter homogeneity assumption does not hold, reporting an average effect of trade openness on the slope of the Phillips curve can be very misleading. Significant effects with different signs can be averaged out while trade openness has indeed played a role in all sample countries. Relaxing the parameter homogeneity assumption, we find that trade openness has significantly changed the slope of the Phillips curve in several major industrial countries.[76]

[76]In our benchmark time series analysis with both trade and financial openness, a significant effect is found in Canada, France, Italy, Sweden and the United States.

Table 4.1: Data sources

Variable	Source
Core CPI inflation	OECD main economic indicators, percentage point
Output gap	OECD economic outlook No. 87, percentage point
Trade openness	Penn World Table, percentage point
Energy price changes	OECD main economic indicators, percentage point
Food price changes	OECD main economic indicators, percentage point
Non-commodity import price changes	OECD economic outlook No. 87, percentage point
Imports	OECD main economic indicators, USD
Nominal GDP	World Development Indicators, current USD
Real GDP	World Development Indicators, constant 2000 USD
Population	World Development Indicators
Financial openness	calculated with data from
	Lane and Milesi-Ferretti (2007), percentage point

Notes: Trade data include both goods and services.
Financial openness is defined as total foreign assets and liabilities divided by GDP.

Table 4.2: Descriptive statistics

	CPI	output gap	openk	openf	energy	food	import	import share	real GDP	population
Australia										
Mean	4.89	-0.26	0.32	1.21	1.62	0.22	-2.68	-0.57	26.46	9.76
Median	3.89	0.04	0.32	1.10	-0.23	0.03	-3.91	-0.78	26.42	9.77
Maximum	11.85	2.16	0.44	2.58	19.74	7.39	15.22	2.71	26.94	9.94
Minimum	-0.22	-4.88	0.22	0.31	-9.31	-5.08	-13.36	-2.80	26.00	9.56
Standard deviation	3.36	1.61	0.08	0.66	6.59	3.22	6.81	1.35	0.29	0.11
Canada										
Mean	4.21	-0.48	0.56	1.48	1.95	-0.39	-1.59	-0.44	27.05	10.24
Median	3.28	0.33	0.54	1.26	1.36	-0.29	-1.90	-0.63	27.02	10.25
Maximum	11.11	3.75	0.75	2.20	20.57	9.88	7.26	1.79	27.49	10.40
Minimum	0.14	-6.20	0.38	0.81	-11.20	-8.59	-10.87	-3.42	26.61	10.07
Standard deviation	3.05	2.41	0.13	0.47	6.03	3.61	4.81	1.37	0.26	0.11
France										
Mean	4.34	-0.02	0.38	2.09	0.60	-0.24	-2.06	-0.46	27.73	10.98
Median	3.10	-0.12	0.35	1.38	0.14	-0.17	-2.20	-0.58	27.73	10.98
Maximum	12.75	2.58	0.56	5.81	15.62	4.99	3.29	0.76	28.04	11.06
Minimum	0.52	-2.44	0.26	0.56	-18.20	-3.53	-8.36	-1.75	27.39	10.90
Standard deviation	3.78	1.41	0.10	1.51	5.94	1.83	3.15	0.69	0.19	0.05
Italy										
Mean	7.46	-0.47	0.43	1.17	-0.77	-1.30	-1.07	-0.20	27.54	10.95
Median	5.50	-0.66	0.43	0.77	-1.07	-0.96	-0.64	-0.16	27.58	10.95
Maximum	22.26	2.46	0.55	2.82	10.75	2.11	8.70	1.58	27.80	10.97
Minimum	1.52	-3.42	0.31	0.43	-15.11	-6.16	-11.57	-2.39	27.18	10.93
Standard deviation	5.98	1.57	0.08	0.76	5.65	2.34	5.01	1.07	0.19	0.01
Japan										
Mean	1.80	-0.16	0.19	0.90	-0.85	-0.83	-2.14	-0.23	28.99	11.72
Median	1.49	-0.32	0.17	0.94	-1.54	-0.67	-1.69	-0.15	29.09	11.73
Maximum	8.61	3.98	0.30	1.94	30.36	2.94	0.84	0.08	29.28	11.76
Minimum	-0.89	-2.85	0.14	0.21	-12.97	-5.48	-9.95	-1.15	28.53	11.64
Standard deviation	2.20	1.81	0.05	0.45	7.16	1.95	2.60	0.31	0.23	0.03

Notes: In this table, CPI, energy, food, import denote core CPI, energy, food, import price inflation, respectively. Openk and openf denote trade and financial openness respectively.
GDP and population are logarithmic values.

Descriptive statistics continued

	CPI	output gap	openk	openf	energy	food	import	import share	real GDP	population
Netherlands										
Mean	3.07	0.08	0.90	3.68	1.49	-1.49	-2.40	-1.28	26.40	9.62
Median	2.59	0.15	0.82	2.65	1.02	-0.99	-2.44	-1.40	26.40	9.62
Maximum	8.44	3.17	1.41	9.73	16.96	5.08	5.53	2.76	26.81	9.72
Minimum	0.53	-3.52	0.60	1.19	-16.52	-6.15	-7.76	-4.11	26.05	9.53
Standard deviation	1.87	1.56	0.25	2.54	7.01	2.26	2.96	1.55	0.25	0.06
Sweden										
Mean	4.81	0.15	0.63	1.97	3.05	-0.38	-0.71	-0.19	26.04	9.06
Median	4.61	0.23	0.54	1.43	3.53	0.26	-1.07	-0.34	26.02	9.07
Maximum	12.47	4.38	0.97	5.12	26.47	5.14	8.25	2.39	26.42	9.11
Minimum	-0.86	-5.80	0.43	0.45	-13.62	-15.89	-12.32	-3.25	25.74	9.01
Standard deviation	4.07	2.46	0.18	1.43	8.09	4.12	4.24	1.30	0.20	0.03
United Kingdom										
Mean	5.43	0.06	0.41	4.29	0.49	-1.09	-2.53	-0.66	27.79	10.97
Median	4.27	0.27	0.38	3.48	0.07	-0.74	-2.15	-0.58	27.77	10.96
Maximum	18.23	4.25	0.60	9.33	13.19	3.70	3.66	1.01	28.20	11.01
Minimum	0.09	-4.02	0.30	1.90	-12.46	-9.54	-13.25	-3.12	27.44	10.94
Standard deviation	4.93	2.01	0.09	2.01	5.71	2.93	3.94	1.02	0.24	0.03
United States										
Mean	4.31	-0.17	0.19	1.03	0.96	-0.61	-2.77	-0.31	29.63	12.46
Median	3.66	0.08	0.18	0.79	0.11	-0.23	-2.35	-0.28	29.62	12.46
Maximum	12.42	2.58	0.29	2.79	21.13	4.28	7.98	0.79	30.07	12.62
Minimum	1.46	-6.16	0.12	0.40	-17.56	-7.03	-13.47	-1.26	29.18	12.30
Standard deviation	2.67	1.86	0.06	0.60	8.60	2.58	4.47	0.46	0.28	0.10

Notes: In this table, CPI, energy, food, import denote core CPI, energy, food, import price inflation, respectively. Openk and openf denote trade and financial openness respectively.
GDP and population are logarithmic values.

Table 4.3: Estimation results of equation (4.3)

Explanatory variable	Australia	Canada	France	Italy	Japan	Netherlands	Sweden	UK	US
Constant	1.38	0.58	0.03	0.12	0.34	0.47	0.21	0.73	0.99
	(0.77)	(0.38)	(0.28)	(0.39)	(0.22)	(0.31)	(0.62)	(0.67)	(0.57)
Lagged core inflation	0.77**	0.90**	0.94**	0.92**	0.66**	0.81**	0.93**	0.81**	0.86**
	(0.12)	(0.08)	(0.04)	(0.05)	(0.23)	(0.09)	(0.10)	(0.09)	(0.11)
Output gap	1.74	-0.82	1.18**	2.12	0.16	0.08	0.94	2.94	0.89**
	(1.20)	(0.48)	(0.42)	(1.40)	(0.49)	(0.29)	(0.69)	(1.79)	(0.30)
Trade openness*output gap	-3.80	-0.89	-2.40	-3.61	0.00	0.22	-0.68	-6.37	-3.69*
	(4.38)	(0.96)	(1.27)	(3.21)	(0.03)	(0.33)	(1.11)	(4.93)	(1.49)
Energy price	-0.00	0.02	-0.01	-0.11	-4.62	-0.00	0.02	0.04	-0.01
	(0.06)	(0.04)	(0.03)	(0.10)	(4.01)	(0.02)	(0.05)	(0.10)	(0.02)
Food price	-0.03	-0.08	0.18	-0.12	0.56	-0.09	-0.19	-0.38	0.10
	(0.15)	(0.11)	(0.10)	(0.13)	(3.32)	(0.06)	(0.12)	(0.22)	(0.09)
Import price	0.05	0.15	-0.21	-0.37	4.23	0.65	0.29	1.04	-0.69*
	(0.51)	(0.33)	(0.44)	(0.30)	(3.04)	(0.52)	(0.74)	(1.26)	(0.33)
Import price*import share	0.11	-0.06	0.82	1.95	-2.13	-1.02	-0.15	-3.45	7.00*
	(2.55)	(1.05)	(1.95)	(1.60)	(3.18)	(0.96)	(2.38)	(4.71)	(3.26)
Adjusted R-squared	0.70	0.86	0.97	0.90	0.84	0.84	0.77	0.80	0.88
Normality	0.59	0.77	0.14	0.00**	0.00**	0.31	0.67	0.06	0.00**
Serial independence	0.89	0.05*	0.36	0.15	0.03*	0.88	0.01**	0.34	0.49
Arch	0.69	0.16	0.70	0.09	0.92	0.17	0.82	0.84	0.87
Structure break	1991**	1987**	1982*			1997*	1992**		1982**

Notes: The dependent variable is the core inflation rate of the respective country.
Standard errors in parentheses. *, ** denotes statistical significance at 5% and 1% level.
Energy price, food price and import price stand for the deviations of the change of these prices from last period's core CPI inflation rate.
Normality: Jarque-Bera test p values.
Serial independence: LM test p values for serial correlation up to two orders.
Arch: test p values for Arch(1).
Structure break: Andrews (1993) Sup test. The number reported is the year with the maximum F statistics.

Table 4.4: Estimation results of equation (4.3) with additional control variables

Explanatory variable	Australia	Canada	France	Italy	Japan	Netherlands	Sweden	UK	US
Constant	2.62**	0.63	-0.29	-0.17	0.18	0.44	-0.98	-1.35**	0.50**
	(0.62)	(0.37)	(0.17)	(0.46)	(0.18)	(0.23)	(0.53)	(0.47)	(0.15)
Lagged core inflation	0.64**	0.46**	0.68**	0.44**	0.75**	0.58**	0.34*	0.09	0.67**
	(0.08)	(0.16)	(0.09)	(0.18)	(0.05)	(0.10)	(0.15)	(0.14)	(0.06)
Lagged core inflation*break dummy	-0.57**	-0.26*							-0.14**
	(0.16)	(0.10)							(0.03)
Output gap	2.12*	1.24**	310.76**	203.43*	-0.21	0.21	1.46**	-0.89	-85.53
	(0.03)	(0.29)	(53.01)	(85.89)	(0.43)	(0.30)	(0.51)	(1.67)	(73.94)
Trade openness*output gap	-5.37	-1.59**	10.24**	11.64	1.63	0.11	-1.41	1.11	-8.55
	(3.43)	(0.56)	(2.30)	(6.92)	(2.33)	(0.32)	(0.81)	(3.67)	(7.86)
Energy price			-0.03*						
			(0.02)						
Food price			0.12*				-0.27**		
			(0.05)				(0.08)		
Import price							0.23**		
							(0.07)		
Import price*import share									
Log population*output gap			-28.63**						-13.27**
			(4.90)						(3.58)
Log GDP*output gap				-7.57*					8.47**
				(3.22)					(2.11)
Trend inflation*output gap								0.23**	0.41**
								(0.06)	(0.12)
Global inflation		0.42**	0.27**	0.71*		0.12*	0.79**	1.20**	0.22**
		(0.14)	(0.09)	(0.32)		(0.06)	(0.18)	(0.19)	(0.05)
Outlier dummy				5.57**	2.34**			-5.03**	2.81**
				(0.78)	(0.68)			(1.27)	(0.43)
Adjusted R-squared	0.80	0.94	0.99	0.98	0.92	0.84	0.88	0.94	0.99
Normality	0.97	0.82	0.22	0.69	0.31	0.25	0.50	0.93	0.77
Serial independence	0.29	0.09	0.80	0.22	0.66	0.84	0.54	0.31	0.39
Arch	0.59	0.07	0.34	0.16	0.89	0.60	0.84	0.68	0.43
Redundancy	0.58	0.76	0.25	0.87	0.19	0.08	0.35	0.74	0.30

Notes: The dependent variable is the core inflation rate of the respective country.
Standard errors in parentheses. *, ** denotes statistical significance at 5% and 1% level.
Energy price, food price and import price stand for the deviations of the change of these prices from last period's core CPI inflation rate.
Normality: Jarque-Bera test p values.
Serial independence: LM test p values for serial correlation up to two orders.
Arch: test p values for Arch(1).
Redundancy: F test p values for redundancy.
Outlier years: Italy (1979-1980), Japan (1980), the UK (1978), and the US (1984).

Table 4.5: Controlling for financial openness

Explanatory variable	Australia	Canada	France	Italy	Japan	Netherlands	Sweden	UK	US
Constant	2.71**	0.64	-0.26	-0.25	0.21	0.54**	-1.04	-1.36**	0.50**
	(0.61)	(0.37)	(0.14)	(0.38)	(0.15)	(0.20)	(0.50)	(0.49)	(0.15)
Lagged core inflation	0.63**	0.41**	0.63**	0.45**	0.69**	0.72**	0.40**	0.08	0.68**
	(0.08)	(0.15)	(0.07)	(0.16)	(0.06)	(0.06)	(0.15)	(0.14)	(0.06)
Lagged core inflation*break dummy	-0.58**	-0.24*							-0.13**
	(0.16)	(0.10)							(0.03)
Output gap	1.45**	1.21**	255.29**	604.71**	-0.46	0.10	5.10*	-0.59	-130.45**
	(0.49)	(0.28)	(49.11)	(158.96)	(0.31)	(0.14)	(1.99)	(0.94)	(21.87)
Trade/financial openness*output gap	-0.72	-0.43**		1.66**	0.47	0.03		0.04	-0.89**
	(0.43)	(0.15)		(0.48)	(0.26)	(0.03)		(0.16)	(0.22)
Trade openness*output gap			14.35**				-10.76*		
			(4.68)				(5.02)		
Financial openness*output gap			-0.51				1.10		
			(0.32)				(0.59)		
Energy price			-0.03*						
			(0.01)						
Food price			0.15**				-0.30**		
			(0.04)				(0.08)		
Import price							0.28**		
							(0.08)		
Import price*import share									
(Log GDP-Log population)*output gap									7.55**
									(1.27)
Log population*output gap			-23.64**						
			(4.56)						
Log GDP*output gap				-23.23**					
				(5.76)					
Trend inflation*output gap				-0.44**				0.22**	0.53**
				(0.13)				(0.06)	(0.08)
Global inflation		0.45**	0.30**	0.74**			0.75**	1.21**	0.21**
		(0.14)	(0.08)	(0.27)			(0.17)	(0.20)	(0.05)
Outlier dummy			1.50**	5.45**	2.94**	2.33**		-5.01**	2.54**
			(0.46)	(0.64)	(0.74)	(0.59)		(1.27)	(0.37)
Adjusted R-squared	0.80	0.94	0.99	0.98	0.92	0.88	0.89	0.94	0.99
Normality	0.98	0.86	0.68	0.41	0.13	0.62	0.85	0.92	0.77
Serial independence	0.38	0.15	0.32	0.28	0.60	0.99	0.78	0.35	0.93
Arch	0.47	0.07	0.23	0.35	0.94	0.96	0.91	0.76	0.48
Redundancy	0.59	0.74	0.10	0.86	0.27	0.50	0.31	0.71	0.20
Restriction	0.67	0.65	0.01**	0.70	0.30	0.79	0.05*	0.81	0.10

Notes: The dependent variable is the core inflation rate of the respective country.
Standard errors in parentheses. *, ** denotes statistical significance at 5% and 1% level.
Energy price, food price and import price stand for the deviations of the change of these prices from last period's core CPI inflation rate.
Normality: Jarque-Bera test p values.
Serial independence: LM test p values for serial correlation up to two orders.
Arch: test p values for Arch(1).
Redundancy: F test p values for redundancy.
Restriction: F test p values for the parameter restriction on financial openness*output gap.
Outlier years: France (1980), Italy (1979-1980), Japan (1980), the Netherlands (1980), the UK (1978), and the US (1984).

Table 4.6: Estimation results with the HP-filtered output gaps

Explanatory variable	Australia	Canada	France	Italy	Japan	Netherlands	Sweden	UK	US
Constant	3.06**	0.85*	-0.33*	0.26	0.12	0.71**	-1.51**	-1.33**	0.87**
	(0.82)	(0.36)	(0.14)	(0.30)	(0.15)	(0.20)	(0.48)	(0.46)	(0.14)
Lagged core inflation	0.57**	0.35*	0.61**	0.56**	0.71**	0.68**	0.22	0.03	0.57**
	(0.10)	(0.15)	(0.07)	(0.16)	(0.06)	(0.05)	(0.13)	(0.14)	(0.06)
Lagged core inflation*break dummy	-0.73**	-0.27**							-0.17**
	(0.21)	(0.10)							(0.03)
Output gap	0.50	1.27**	375.61**	719.03**	-0.34	0.18	-2.40	0.18	-144.24**
	(0.60)	(0.26)	(63.60)	(159.61)	(0.26)	(0.12)	(1.62)	(1.31)	(25.24)
Trade and financial openness*output gap	-0.09	-0.42**		1.93**	0.37	0.01	0.63	-0.14	-1.16**
	(0.47)	(0.14)		(0.45)	(0.24)	(0.02)	(0.39)	(0.24)	(0.35)
Trade openness*output gap			25.78**						
			(5.21)						
Financial openness*output gap			-1.23**						
			(0.35)						
Energy price			-0.04**						
			(0.01)						
Food price			0.22**				-0.20**		
			(0.05)				(0.07)		
Import price							0.24**		
							(0.06)		
Import price*import share									
(Log GDP-Log population)*output gap									8.36**
									(1.47)
Log population*output gap			-34.89**						
			(5.90)						
Log GDP*output gap				-26.06**					
				(5.77)					
Trend inflation*output gap				-0.53**			0.34*	0.19**	0.57**
				(0.15)			(0.15)	(0.07)	(0.08)
Global inflation		0.42**	0.35**	0.49*			1.00**	1.24**	0.22**
		(0.14)	(0.08)	(0.26)			(0.16)	(0.20)	(0.05)
Outlier dummy			1.42**	5.82**	2.75**	2.24**		-4.40**	2.30**
			(0.48)	(0.57)	(0.71)	(0.58)		(1.32)	(0.39)
Adjusted R-squared	0.70	0.94	0.99	0.98	0.92	0.88	0.91	0.93	0.99
Normality	0.85	0.96	0.60	0.15	0.07	0.66	0.42	0.81	0.17
Serial independence	0.08	0.39	0.30	0.68	0.63	0.98	0.24	0.39	0.98
Arch	0.76	0.27	0.27	0.44	0.97	0.77	0.21	0.70	0.41
Redundancy	0.17	0.82	0.17	0.54	0.13	0.33	0.20	0.66	0.29
Restriction	0.34	0.38	0.00**	0.34	0.62	0.92	0.24	0.96	0.36

Notes: The dependent variable is the core inflation rate of the respective country.
Standard errors in parentheses. *, ** denotes statistical significance at 5% and 1% level.
Energy price, food price and import price stand for the deviations of the change of these prices from last period's core CPI inflation rate.
Normality: Jarque-Bera test p values.
Serial independence: LM test p values for serial correlation up to two orders.
Arch: test p values for Arch(1).
Redundancy: F test p values for redundancy.
Restriction: F test p values for the parameter restriction on financial openness*output gap.
Outlier years: France (1980), Italy (1979-1980), Japan (1980), the Netherlands (1980), the UK (1978), and the US (1984).

Table 4.7: Seemingly unrelated regressions results

Explanatory variable	Australia	Canada	France	Italy	Japan	Netherlands	Sweden	UK	US
Constant	0.12*	-0.03	-0.05*	-0.07	0.02	0.01	-0.07	-0.08*	0.02
	(0.06)	(0.03)	(0.02)	(0.07)	(0.02)	(0.03)	(0.05)	(0.04)	(0.02)
Lagged core inflation	0.80**	0.11	0.54**	0.02	0.79**	0.75**	0.47**	0.21**	0.52**
	(0.09)	(0.10)	(0.10)	(0.29)	(0.08)	(0.10)	(0.13)	(0.12)	(0.10)
Lagged core inflation*break dummy	-0.30	-0.07							-0.43**
	(0.48)	(0.13)							(0.10)
Output gap	1.13**	1.01**	342.81**	670.38**	0.20	0.41**	0.96**	-164.93	-267.04**
	(0.49)	(0.22)	(43.14)	(317.13)	(0.30)	(0.14)	(0.22)	(88.32)	(39.07)
Trade and financial openness*output gap	-0.52	-0.33**		1.36	-0.07	-0.01	-0.15*	-0.36	-1.86**
	(0.44)	(0.12)		(0.92)	(0.25)	(0.02)	(0.07)	(0.19)	(0.38)
Trade openness*output gap			20.08**						
			(4.24)						
Financial openness*output gap			-0.78**						
			(0.30)						
Energy price			-0.02*						
			(0.01)						
Food price			0.16**				-0.23**		
			(0.04)				(0.07)		
Import price							0.21**		
							(0.06)		
Import price*import share									
(Log GDP-Log population)*output gap									15.47**
									(2.27)
Log population*output gap			-31.76**						
			(4.00)						
Log GDP*output gap				-24.29*					
				(11.49)					
Trend inflation*output gap				-0.50*				0.33**	0.92**
				(0.26)				(0.10)	(0.11)
Global inflation		0.80**	0.42**	1.46**		0.11**	0.59**	0.87**	0.33**
		(0.14)	(0.10)	(0.47)		(0.05)	(0.15)	(0.20)	(0.09)
Outlier dummy			-0.08*	0.33	0.03	-0.04		-0.03	-0.06
			(0.04)	(0.22)	(0.08)	(0.07)		(0.10)	(0.05)
Adjusted R-squared	0.76	0.95	0.99	0.93	0.83	0.85	0.88	0.95	0.97

Notes: The dependent variable is the core inflation rate of the respective country.
Standard errors in parentheses. *, ** denotes statistical significance at 5% and 1% level.
Energy price, food price and import price stand for the deviations of the change of these prices from last period's core CPI inflation rate.
Outlier years: France (1980), Italy (1979-1980), Japan (1980), the Netherlands (1980), the UK (1978), and the US (1984).

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